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An examination of the evolving relationship between interest rates of different maturities in Japan, and test of the expectations hypothesis of the term structure to ascertain the feasibility of using asymmetric dynamics in yield spreads

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An examination of the evolving relationship between interest rates of different maturities in Japan, and test of the expectations hypothesis of the term structure to ascertain the feasibility of using asymmetric dynamics in yield spreads

by

Shew-Huei Kuo

A dissertation submitted to the graduate faculty

in partial fulfillment of the requirements for the degree of

DOCTOR OF PHILOSOPHY

Major: Economics

Major Professor: Walter Enders

Iowa State University

Ames, Iowa

2000

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For the Graduate College

to my family

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ABSTRACT

The purpose of the study was to examine the evolving relationship between interest rates of different maturities in Japan, allowing for possible asymmetrical movements toward a long-run equilibrium rather than use of the traditional analysis of symmetrical movements. The relationship among interest rates of different maturities has been extensively studied with mixed results. This research provides empirical evidence concerning the properties of interest rate term structure using Japan's data from 1980 to 1998. The study employed the testing methodologies which permit asymmetry in the adjustment toward equilibrium in the threshold autoregressive and the momentum-threshold autoregressive specifications. The test results support the expectations hypothesis of the term structure of interest rate. It was shown that the error correction toward the long-run equilibrium relationship is best estimated as an asymmetric process.

CHAPTER 1. INTRODUCTION

The past decade has seen numerous empirical studies that relate to the expectations hypotheses of the term structure. Specifically, the studies examine the spread between interest rates of different maturities and explore to what extent the yield curve can be applied to predict the future movements of interest rates. The expectations hypothesis, in its pure form, maintains a zero term premium and the same expected rate of return over the next holding period for bonds of all maturities. Securities of different maturities are perfect substitutes under the assumption of risk neutrality. Thus, for a given holding period, a long-term rate is an unbiased average of current and expected future short-term rates. A rise in the long-term rate relative to the short-term rate reflects the expectation of higher short-term rates in the future. Given that market predictions are unbiased and correct, on average, subsequent short-term rates would be apt to rise. The restriction of a zero premium is not posited in the general form of the hypothesis. Other than interest rate expectations alone, the term structure is determined by a consideration of liquidity and maturity preferences.

Campbell and Shiller (1984), Mankiw and Summers (1984), and Mankiw (1986) observed the subsequent change of the long-term rate that is inconsistent with the direction predicted by the spread. On the other hand, Campbell and Shiller (1987), Hardouvelis (1988), and Mishkin (1988, 1990a, 1991) documented increases in future short rates when the spread arises. Recently, Campbell and Shiller (1987) showed that the expectations hypothesis necessitates a cointegrating relationship between short- and long-term rates. Using tests of cointegration, they studied the predictive power of the term structure in forecasting expected future short-term rate changes, and found strong support for the expectations hypothesis.

The usual unit-root tests and cointegration tests are conducted in the context of the Dickey and Fuller (1979), Johansen (1996), Engle and Granger (1987), and Stock and Watson (1988) methodologies, that assume a linear symmetric adjustment process. However, empirical evidence often suggests that cointegrated macroeconomic variables display asymmetric adjustment toward a long-run equilibrium. The concept of asymmetric behavior in business cycles is discussed in Keynes (1936), which states that contractions in an economy are often more violent but swifter than expansions. In the asymmetric price adjustment model of Ball and Mankiw (1992), shocks that raise firms' desired price trigger larger price responses than shocks that lower desired prices. The degree of asymmetry that effects of monetary shocks have on output is positively correlated with the expected inflation in the analysis of postwar U.S. data of Rhee and Rich (1995). Tinsley and Krieger (1997) reported that shortfalls of production from trend are larger than positive deviations, while the increases of price levels take place more readily than the decreases. Teräsvirta and Anderson (1992) estimated the smooth transition autoregressive model of the production index for 13 OECD countries, and found stronger responses of industrial production to large negative shocks associated with oil crises than to positive shocks. Granger and Lee (1989) showed the asymmetric adjustment process in a cointegrated series of U.S. production, sales, and inventories. Asymmetric adjustments are also characterized as deep versus steep movements by Sichel (1993). He provided evidence indicating deepness in U.S. unemployment, industrial production, and GNP, whereas evidence of sharpness was found only in unemployment.

It has also been argued that the dynamic adjustments of yield spreads might be different in the presence of positive versus negative deviations from a long run equilibrium. The behavior of the term structure may incorporate an inherently asymmetric process in inflation rates and the business cycle through the comovement of nominal interest rates and various macroeconomic variables. The asymmetry of yield spreads may also arise from an asymmetric policy response of the monetary authority to inflationary pressures and unemployment levels, as well as exchange rate movements. Mankiw and Miron (1986) argued that the manner in which the Federal Reserve controls interest rates affects the prediction of the term structure for future short rate changes and the validity of the test of the rational expectations hypothesis. Rudebusch (1995) provided an explanation to the varying predictive content of the yield curve using a model that relates the term structure to the behavior of the Federal Reserve. The asymmetric response of ex-ante real and nominal interest rates to money innovations was found by Guirauis (1994). Enders and Siklos (1998), and Enders and Granger (1998) provided empirical evidence of asymmetric adjustments in short-term and long-term rates in the analysis of various U.S. Treasury yields.

Under the monetary transmission mechanism, monetary actions which operate directly on short-term rates change the whole spectrum of the yield curve. On this basis, the term structure can be defined as anticipations of future short rates and term premium. Nevertheless, Japanese money markets for securities at different horizons, such as long-term and short-term bond markets, were segmented, and the markets were small-scale, illiquid, and strictly administered. Markets for different varieties of securities were not soundly linked and, accordingly, yield spreads did not contain information on interest rate expectations or classically-defined risk.

Yet, over the past 15 years, when the financial markets in Japan were expeditiously liberalized, money markets began to expand and become crucial to Japanese corporate finance, which made the expectations theory of term structure applicable to the analysis of Japanese interest rate behavior.

The implicit assumption of symmetric adjustment is problematic, since empirical evidence often indicates that many macroeconomic variables display asymmetric adjustment paths. Nevertheless, asymmetries in term structure have seldom been formalized and accounted for in tests of unit-root and cointegration. The purpose of the present study was to examine the relationship between interest rates of different maturities in Japan, allowing for possible asymmetrical movements toward a long-run equilibrium rather than use of the traditional analysis of symmetrical movements. Sickel (1993) discussed two types of asymmetric cycles in time series: deepness and steepness. Deepness occurs when troughs are more pronounced that peaks and steepness occurs when contractions are steeper than expansion. Corresponding to these two versions of asymmetry, this research proposes asymmetric test methodologies in the frameworks of Enders and Granger (1998) and Enders and Siklos (1998), to account for threshold autoregressive and momentum threshold autoregressive adjustments in the term structure. This analysis of term structure would avoid a misspecified cointegration and enhance the power as well as the size property of tests in the presence of asymmetric adjustment.

The organization of this research is as follows. Chapter 2 reviews the theoretical basis of the expectations hypothesis of interest rate term structure, the evolution of empirical tests, and the results of the tests. Chapter 3 presents further theoretical arguments for the information content of term structure. The last section of Chapter 3 stresses on the potential role of monetary policy in interpreting the predictive power of term structure for future short rate changes ,and in influencing the validity of the test of the rational expectation hypothesis.

The explanations of term structure asymmetry which are consistent with the expectations hypothesis are emphasized in Chapter 4. Three possible sources for asymmetric adjustment in term structure are carefully discussed to justify the application of asymmetric unit-root and cointegration tests in Chapter 7. The operation instruments and procedures of the Bank of Japan, and the phases of monetary policy experience of Japan are discussed in Chapter 5.

Chapter 6 lays out the framework for analysis of the term structure of interest rates. A detailed description of prevailing test methodologies as well as newly developed asymmetric test methodologies is presented. It first covers conventional unit-root and cointegration tests, which relate specifically to Dickey-Fuller, Engle-Granger, and Johansen methodologies. The following section then introduces the newly-advanced Enders-Granger unit-root and Enders-Siklos cointegration tests that enable asymmetric dynamic process.

Chapter 7 discusses data sources and studies the behavior of term spreads by employing the new asymmetric unit-root and cointegration tests to account for the essence of asymmetric adjustments in interest rate term structure. Detailed interpretation of the results are given. To provide further insight into the dynamic adjustment process of the spread, asymmetric error-correction models are estimated in the last section of this chapter. Finally, Chapter 8 summarizes the findings and recommendations for further study are given.

CHAPTER 2. THEORETICAL FRAMEWORK

Theory of Term Structure

The term structure of interest rates measures the relationship of yields on default-free pure discount instruments of different term maturities. The literature on the term structure has centered around four strands of thought: (a) investor expectations, (b) liquidity preference, (c) market segmentation, and (d) equilibrium term structure.

There are several versions of the expectations hypothesis. The pure expectation hypothesis (PEH), advanced by Irving Fisher (1930), is based on the following assumptions: (a) the same market expectation of future interest rates is formed by each investor, and (b) only a security sequence that yields the highest expected holding period return will be chosen by investors. As a result of maturity arbitrage among investors, it is asserted that the term premium is zero, forward rates are unbiased estimates of expected future spot rates, and rational investors may either invest in a k-period bond or roll over a series of one-period bonds for k successive periods. Various forms of the pure expectations hypothesis are characterized by the time horizon over which the excess returns on long-term over short-term instrument are zero. The one-period form states that the expected one-period returns on one-period and n-period bonds are equivalent. The n-period form states that the expected n-period returns on one-period and n-period bonds are equivalent. However, when interest rates are random, Jensen's inequality (cited in Campbell, Lo, & MacKinlay, 1997) implies that the pure expectation hypothesis in its one-period form does not agree with its n-period form. It can be shown that, in the lognormal homeskedastic model of interest rates, the expected excess one-period log return on an n-period bond over a one-period bond, derived from the one-period PEH equation, will not be the same as that derived from the n-period PEH equation. The log pure expectations hypothesis expresses the PEH in log forms. The expected excess one-period log return on an n-period over a one-period bond is assumed to be zero under this form.

In its log form, the pure expectations hypothesis imposes the restriction that the term premium is zero, while a more general version of this hypothesis, called the strong form of the expectation hypothesis (EH), states that the term premium is constant over time (Cook & Hahn, 1990; Murphy, 1986; Russell, 1992). For lognormal homoskedastic bond yields, the EH holds in its one-period, n-period, and log form. The consistency of the form is due to the time-independence of the Jensen's inequality effects.

The pure expectations hypothesis assumes investors to be risk neutral. The liquidity preference hypothesis (Hicks,1946; Lutz, 1940), on the contrary, states that investors are influenced by the uncertainty. The assertion is that price volatility increases with duration, thus the liquidity premium is the increment needed to compensate holders of longer-term bonds for bearing additional capital risk. The return on long-term bonds should be systematically greater than those on short-term bonds, and should also increase with maturity. On the other hand, Culbertson (1957) articulated the market segmentation hypothesis, based on the assumption that investors have strong maturity preferences and deal only in the spot market security that corresponds to their holding period. There has been limited arbitrage across the maturity spectrum. Securities of different maturities trade in separate markets, and the demand and supply in each separate market determines security returns. Modigliani and Sutch (1966, 1967) combined some features of the market segmentation hypothesis with the expectations model in their preferred habitat hypothesis. They postulated that investors prefer holding bonds with maturities matching their investment planning horizon; thus, the risk premium is the increment required to induce investors to mismatch their planning horizon and the time until maturity. This approach allows for heterogeneous preferred habitats and leads to any possible pattern of term premium, with no restriction on the sign or monotonicity.

Another version which expresses the expectations hypothesis in the weak form assumes a time-varying term premium (Shiller, Campbell, & Schoenholz, 1983; Fama, 1984a; Mankiw & Summers, 1984; Mankiw, 1986; Mankiw & Miron, 1986). The assertion is justified on the ground that the term premium is likely to change in response to interest rate uncertainty induced by the changes in fiscal and monetary policy, interest rate level, and cyclical factors (Dua, 1991).

The empirical results in Mankiw and Miron (1986), Cook and Hahn (1990), Dua (1991), and Fama (1990) show little evidence in support of the pure expectations hypothesis. Limited support of the expectations hypothesis in general form is found in relatively earlier literature. Mankiw (1986) incorporated changes in perceived risk and relative asset supplies to the description of the large variation in the term premium. His investigation of the term structure of U.S., Canada, the U.K., and Germany reveals the failure of the expectations hypothesis to explain observed fluctuations in interest rates. The analyses of U.S. postwar data by Campbell and Shiller (1984), and Fama (1984) reject the expectations hypothesis. Nonetheless, Froot (1989) used survey data of interest rate expectations in the U.S., and found support for the expectations hypothesis at long maturities.

The Expectations Hypothesis and Cointegration

In their present value model, Campbell and Shiller (1987) showed that the expectations hypothesis necessitates a cointegrating relationship between short- and long-term rates. In the literature, numerous empirical studies on the expectations hypothesis have focused on tests of cointegration and the predictive power of term spread in the forecast of expected future short rate changes.

Traditional research in modeling economic time series often employs univariate, autoregressive, integrated moving average representations. Each univariate economic series is expressed as containing a stochastic trend, which involves a unit root in the autoregressive process. However, equilibrium theories and empirical evidence imply that stochastic trends that characterize the processes might be the same among several economic variables, thus the vector of stochastically trending variables has a reduced rank. The concept of cointegration was formulated by Engle and Granger (1987). An n-dimensional nonstationary series X_t which can be transformed to yield a stationary invertible ARMA representation after differencing d times is said to be integrated of order d, denoted by $X_t \sim I(d)$. Given that all components of the vector series $X_t=(x_{1t},x_{2t},\ldots,x_{nt})'$ are integrated of order d, there may exist linearly independent vectors $\beta_i=(\beta^i_1,\beta^i_2,\ldots,\beta^i_n)'$, i = 1,...,r, such that the linear combination $\beta_i'X_t= \beta^i_1x_{1t}+\beta^i_2x_{2t}+\ldots+\beta^i_nx_{nt}$ is integrated of order (d-b), b>0, denoted by $X_t \sim CI\ (d,b)$, with the cointegrating rank r. The number of linearly independent cointegrating vectors β_i's that span the cointegrating space can be at most n-1 for n-dimensional series X_t.

For a vector series X_t that is cointegrated of order (1,1) with cointegrating vector β_i, i=1,...,r, X_t achieves stationarity after first differencing, but there are r linear combinations of X_t

such that $\beta' X_t$ is stationary, where $\beta = (\beta_1, \beta_2, \ldots, \beta_r)$. $\beta' X_t = 0$ can be interpreted as the long run equilibrium. The cointegration implies that the deviation from long-run equilibrium, denoted as $e_t = \beta X_t$, is stationary. Thus, the deviation is temporary in essence and is called equilibrium error.

The integration and cointegration within the term structure as explained by Hall, Anderson, and Granger (1992), are shown by the following. The essence of the Fisher-Hick model can be written as:

$$R_{k,t} = \frac{1}{k} \sum_{j=1}^{k} F_{j,t} \quad , \text{ for } k = 1, 2, 3, \ldots \tag{2.1}$$

where $R_{k,t}$ denotes the time t log compounded yield to maturity on a k period pure discount bond, and $F_{k,t}$ denotes the time t log return on the time t+k-1 investment of a one-period bond which matures at time t+k.

The (j-1)-period-ahead one-period log forward rate is assumed to be equal to the expected one-period log bond rate j-1 periods ahead, plus the term premium:

$$F_{j,t} = E_t(R_{1,t+j-1}) + \Lambda_{j,t} \tag{2.2}$$

where E_t denotes the expectations operator, conditioned on information available at time t, and $\Lambda_{j,t}$ denotes the term premium.

Equations (2.1) and (2.2) lead to the general statement of expectations hypothesis of the form:

$$R_{k,t} = \frac{1}{k} \left[\sum_{j=1}^{k} E_t(R_{1,t+j-1}) \right] + L_{k,t} \tag{2.3}$$

where $L_{k,t} = \sum_{j=1}^{k} \Lambda_{j,t}$ denotes the k-period term premium. The log yield to maturity on a k-period pure discount bond equals the average of the current and expected future one-period log return for k successive periods, plus a term premium.

The cointegration between yields of different maturities is implied in equation (2.3) as:

$$S_{(k,1),t} \equiv R_{k,t} - R_{1,t} = \frac{1}{k}\sum_{i=1}^{k-1}\sum_{j=1}^{j=i} E_t \Delta R_{1,t+j} + L_{k,t}$$

where $\Delta R_{k,s} = R_{k,s} - R_{k,s-1}$ (2.4)

If $\Delta R_{1,t}$ and the premium $L_{k,t}$ are stationary, then the yield spread $S_{(k,1),t} = R_{k,t} - R_{1,t}$ is also stationary. Given that yields to maturity are integrated of order 1 as suggested by causal observation, the expectations hypothesis states that there should be a cointegrating vector of the form $(1, -1)'$ for $[R_{k,t}, R_{1,t}]'$. $R_{k,t}$ is cointegrated with $R_{1,t}$. Thus, for an n-dimensional yield series $[R_{k1,t}, R_{k2,t}, \ldots, R_{kn,t}]'$, there are n-1 linearly independent n-dimensional spread vectors that span the cointegrating space: $[(-1, 1, 0, \ldots, 0)', (-1, 0, 1, 0, \ldots, 0)', \ldots, (-1, 0, \ldots, 0, 1)']$. The long-run equilibrium, which is defined in Brennen and Schwartz (1979), Cox, Ingersoll, and Ross (1985), Richard (1978), and Vasick (1979), is the condition where the yields and premium approach constant levels such that $\Delta R_{1,t+j} = 0$ and $L_{k,t} = \beta_0$. β_0 is a constant equilibrium value and is expressed as $\beta_0 = R_{k,t} - R_{1,t}$.

The above statement can be generalized to imply that yields of different maturities can be modeled as a cointegrated system. Hall, Anderson, and Granger (1992) showed that a cointegrating relationship exists between any two yields of the n series. The n-1 linearly independent cointegrating vectors, which correspond to the spread vectors for one-period and k-period yield, constitute the basis of the cointegrating space for any set of n yields $[R_{k1,t}, \ldots, R_{kn,t}]'$.

An alternative analysis of the cointegration and the error correction is provided by Stock and Watson (1988). It is shown that an n-dimensional cointegrated process X_t with a cointegrating order of (1,1) and n-p linearly independently cointegrating vectors, can be represented in terms of a linear combination of p common stochastic trends plus a stationary component. Under the expectations hypothesis, there exists at most one common stochastic trend underlying each yield of different maturities. That a stochastic trend in the short-term yield will also be passed onto the long-term yield is implied by the equivalence of the return on holding an n-period instrument to maturity and the expected return on repeated investment in a series of a one-period instrument. The yield curves of different maturities are linked and shaped by the nonstationary common factor, thus the yield series could be modeled as a cointegrated system.

Campbell and Shiller (1991) provided another insight into the error correction representation of the term structure. In the context of the rational expectations present value model, the spread between long- and short-term rates is an optimal forecast of a weighted sum of future changes in short-term rates, contingent on the full information set of the market participants. Therefore, the spreads reflect the superior market information used by agents apart from the current and lagged changes in short-term rates, and agents' forward-looking behavior is accounted for by the error correction mechanism.

The representation theorem advanced by Granger (1983) shows that cointegrated series can be represented by error correction mechanisms, which relate the changes in variables to past deviations from the long-run equilibrium and to past changes in variables of cointegrated system. Engle and Granger (1987) proposed a consistent and efficient two-step estimator for cointegration and error correction model. The estimation involves least square regressions of the cointegration relationship and the error correction structure, in which parameter estimates of the cointegration vectors at the first regression are used in the error correction form at the second stage. It is indicated in their empirical analysis that the U.S. short- and long-term rates are cointegrated over the 1952-1982 sample period. Campbell and Shiller (1987) showed that the yields to maturity of U.S. bonds, over the 1959-1983 sample period, are cointegrated. They evaluated the variance between the actual and the theoretical spread predicted by the VAR model, and found support for the rational expectations hypothesis that yield spread defines the cointegrating vectors and provides an optimal forecast of the present value of future changes in short-term rates. Stock and Watson (1988) proposed tests for a reduced number of common trends in a multivariate time series. The tests were developed based on the implications of the Granger representation theorem and the correction of the OLS first-order autoregressive matrix. They applied the tests to U.S. postwar data and found that the interest rates are cointegrated and share at least one common trend that drives the long-run behavior of the individual series. Bradley and Lumpkin (1992) adopted the two-step procedure of Granger and Engle and found a cointegrated system among the U.S. Treasury yield curve for securities ranging in maturity from three months to 30 years.

MacDonald and Speight (1988) adopted a cointegration methodology for a bivariate autoregressive (BVAR) system, which consisted of the yield spread and the changes in short rates, to take account of the restrictions implied by the expectations hypothesis and the further assertions of the hypothesis that spreads Granger-cause changes in short-term rates. The expectation hypothesis was found to be supported by U.K. data. MacDonald and Speight (1991) extended the analysis to allow for a time-varying term premium, additional interest rate volatility test, and a broader information set of macroeconomic variables for the orthogonality test. For their studies of U.K. and U.S. data, the expectations hypothesis could not be rejected under certain conditions, while with Belgium, Canada, and Germany data, the hypothesis was strongly rejected. A similar cointegration technique in a BVAR system was employed by McFadyen, Pickerill, and Devaney (1991). Their results based on the yields on U.S. Treasury issues from 1953 to 1984 conclude that long-term rates Granger-cause changes in short-term rates. Mandeno and Giles (1995) applied a bootstrap simulation procedure to obtain modified critical values for the augmented Dickey-Fuller cointegration test which account for precise forms of structural breaks in the data. Their study on the U.S. Treasury securities over the 1950-1982 sample period yielded strong evidence of a causal relationship from long- to short-term rates, and some evidence of causality in the reverse direction. The same type of technique was used in the analysis of Australian data by Karfakis and Moschos (1993), and some support for the expectations hypothesis was obtained. However, the analysis using similar methodology by Shea (1992) yielded results less favorable to the expectation hypothesis with U.S. data.

CHAPTER 3. INFORMATIONAL CONTENT RELATED TO TERM STRUCTURE

The determination of interest rates in money and capital markets play an important role in the allocation of economic resources. The rates influence and are influenced by savings, investment, and the share of portfolio wealth. Oftentimes in the framework of monetary policy, the term structure of interest rates is employed as an indicator of market expectations of the policy stance. It is mainly recognized to possess informational and predictive content concerning inflation and real activity, which might be valuable to the monetary authority in policy design and implementation. Market participants, in practice, also pay relative attention to the stance of monetary policy when forecasting future rates. In the context of models of nominal term structure, monetary authority influences interest rates generally through inflation expectations, and through short-term liquidity effects as well. A number of studies investigated the relationship between the instruments of monetary policy and the yield spread between long- and short-term interest rates. The policy actions, in most cases, change in response to changes in the information content of the term structure. The monetary authority frequently intervenes into the money markets and uses short-term rates as the primary policy instruments. Hence, it is often suggested that the yield curve may reflect useful information with regard to the stance of monetary policy and, to some extent, is influenced by the central bank.

The theoretical basis for the information content of term structure with respect to market expectations of future movements in real activity and inflation, is that the nominal interest rate of a given maturity consists of the ex ante real interest rate, expected inflation, and term premium. The relationship between the yield curve and the three above mentioned components is complex over the time horizon of the underlying security. The information in the yield curve for forecasting future changes in short rates varies with the maturity of the rates involved. The next three sections discuss theoretical argument and empirical research on the information in the term structure of interest rates.

Instruments of Monetary Policy and their Effects on Term Structure

Yield spread reflects the stance of monetary policy. As maintained by the expectations theory of the term structure of interest rates, long-term yields are the weighted average of current and expected future short-term yields plus a term premium. Monetary policy affects long-term yields to the extent that it influences current short-term yields and alters the market expectations of future short-term yields.

The market for reserves held by depository institutions reveals the stance of monetary policy. The operating procedure of a central bank mainly involves the control of the reserves of the banking system. Bank reserves can be accurately managed by being accounted as a liability entry in the balance sheet of the central bank. Changes in bank reserves affect the capacity for credit expansion in banks, and thus they revise short-term interest rates and growth in monetary aggregates. The central bank can alter the aggregate supply of bank reserves through discount window lending and open market purchases or sales of Treasury securities. Providing fewer reserves than demanded by banks leads to rises in interest rates prevailing in interbank reserve markets, while supplying more reserves than banks demand leads to a decline in interbank rates.

Short-term yields move closely with the interest rate that serves as an instrument of monetary policy. Since short-term borrowing acts as a reasonably close substitute for interbank borrowing, an increase in the rates of interbank reserves should be accompanied by an increase in other short-term interest rates. Similarly, because long-term rates are linked to the current and expected future path of short-term interest rates, current expectations of future policy moves of the central bank play a crucial role in the movements of long-term rates.

In their analysis of policy scenarios, Roley and Sellon (1995) concluded that the path and magnitude in which long-term rates move in response to policy actions depend on the investors' views on the probability of future policy actions and the expected persistence of policy actions. The analysis separates from alternative conditions under which interest rates are affected by variation of real interest rates or inflationary expectations.

The scenario in which a policy action is expected to persist over the entire investment horizon indicates a positive reaction of short-term as well as long-term rates to the policy action implied by the monetary transmission mechanism. If a policy action is interpreted as the first stage in a sequence of further reinforcement of actions, movements in medium- and long-term rates will fully or even exceedingly reflect the current policy move. Nevertheless, in the scenario that a current policy action is viewed as merely transitory and expected to be offset or eventually reversed hereafter, long-term rates show little or an inverse response to the monetary policy action in comparison with short-term rates.

Rudebusch (1995) estimated a model of Federal Reserve interest rate targeting behavior, which, in conjunction with the hypothesis of rational expectations, illustrates the link between Fed behavior and the predictive content of the term structure. In the analysis by Rudebusch, the information described by yield spreads between an overnight Fed fund rate and one-month or three-month rates indicates the temporary daily fluctuations from the persistent target rate. The subsequent daily rates are expected to move back to the target level. Therefore, the prevailing three-month rate is close to the target rate. Yield spreads between short-term bills, such as 30-

day and 60-day bills, account for the progressive essence of policy actions. The Fed more often adopts sequential target adjustments to attain a required sizable target change. The anticipated changes in interest rates over the period between the inauguration and the completion of policy actions are embodied in the yield spreads. Yield spreads containing medium-term bills with maturities between three and twelve months are correlated with the target persistence, as the foreseen near-term target change is included in current longer-term rates. Spreads concerning long-term bonds appear to reflect the Fed's expected and actual containment of inflation, with the maintained stationarity of real interest rate. The manner in which the Federal Reserve manages the Fed fund rate accounts for the varying predictive power of yield spreads for future interest rates at various time horizons. The monetary policy actions are reflected in the yield spread.

Yield Curve as a Predictor of Future Inflation

In the theoretical framework of credit market theory, interest rates are determined by the supplies and demands for credit arisen from surpluses and deficits of the major sectors of the economy. Accordingly, the yield spread responds not only to the policy-driven shifts in short-term rates but also to the shifts in long-term rates that reflect the state of the economy and thus the conditions on credit markets. Based on the assumption that a change in long-term yields is caused by a change in demand for credit among sectors, the standard view of credit market theory predicts that the change in the yield spread will precede fluctuations in real economic activity and inflation.

An alternative framework for the analysis of the ability of the yield spread to predict a change in the future inflation rate consists of a combination of the Fisher equation and the rational expectations hypothesis of term structure. The first component of the joint hypothesis, introduced by Irving Fisher (1930), maintains that the nominal interest rate of a given maturity is the sum of the ex ante real interest rate and the expected inflation for the period from the present to the maturity of the instrument. The second component of the joint hypothesis states that the investors' forecasts of the future path of the interest rate implicit in the term structure are optional or unbiased forecasts, and arbitrage ensures variation in the investors' forecasts of future rates. The forecast is fully reflected in the variation in interest rates of various maturities. After adjusting for the liquidity or risk premia, the expected 1-period return from a bond of n-periods to maturity is the same as the certain return from a 1-period bond. By combining these relationships, the following expression can be obtained (Tzavalis & Wickens, 1997).

$$R_{k,t}=E_t r_{k,t}+E_t\Pi_{k,t}+\varphi_{k,t} \tag{3.1}$$

where $R_{k,t}$ = compounded yield to maturity on a k-period bond at time t,

E_t = expectations operator, conditioned on information available at time t,

$r_{k,t}$ = average real interest rate from time t to t+k,

$\Pi_{k,t}$ = average inflation rate from time t to t to t+k,

$\varphi_{k,t}$ = term premium on a k-period bond till maturity,

all components are in the form of logarithm, except for the inflation rate, which is expressed as the first difference of two logarithms.

Equation (3.1) (above) for the 1-period rate can be subtracted from equation (3.1) for the k-period rate t to obtain

$$R_{k,t}-R_{1,t}=E_t[r_{k,t}-r_{1,t}]+E_t[\Pi_{k,t}-\Pi_{1,t}]+\varphi_{k,t}-\varphi_{1,t} \tag{3.2}$$

Equation (3.2) implies that yield curve spreads contain information on changes in ex ante real interest rate, market forecasts of future path of the inflation rate, and term premia. Several research studies in the literature suggest the use of an intertemporal general equilibrium model to study the term structure of interest rates. In those dynamic macroeconomic models, equation (3.2) is conventionally included in a way that is fully consistent with maximizing behavior and rational expectations, and represented as a semi-reduced form of interest rates.

Under the assumptions of rational expectations, the time-invariance of ex ante real rate and the constancy of term premia, the nominal term structure solely reflects market's forecasts of future inflation. Fama's (1975) classic study on interest rates accepted the joint hypothesis that the real rate of interest is constant and that expected inflation is an unbiased prediction of the actual inflation. Subsequent work (Nelson & Schwert, 1977; Huizinga & Mishkin, 1986; Mishkin, 1990a, 1990b, 1990c, 1991, 1992) concluded that term structure has explanatory power for predicting changes in inflation, as the test results reject the null hypothesis that the coefficient on the yield spread, which reveals information content of future inflation path, is zero. Nevertheless, the hypothesis of the constancy of real rate, under which the test equation is derived, is in controversy, since the results also refject the null hypothesis that the coefficient on term spread is one in the same equation.

The empirical analyses of Mishkin (1990b) and Jorion and Mishkin (1991) indicate that long-term interest rates reflect expected inflation more fully than do short-term rates. These results are explained analytically by the differences in the relative variability as well as the correlation of expected future inflation change with real term structure slope. A relatively large variability and a negative correlation of real rate changes to the expected inflation weaken the predictive power of yield curve spreads in forecasting future changes in inflation.

Many studies have been conducted on the issue of the questionableness of the claim that the real rate of interest is literally constant. It is suggested that the variability of the real interest rate is the greatest in the short run, while in the absence of future disturbance, the variability is less and the rate will converge to a constant in the long run. As the expected inflation will be embodied in the nominal interest rate progressively with the passage of time, the term structure at longer maturities will contain significant predictive power in the forecasting future changes in inflation.

In another framework that allows price stickiness and short-term variations in real interest rate, Frankel (1982) and Frankel and Lown (1994) derived an alternative hypothesis to the null hypothesis rejected by Mishkin (1992). The interest rate of a given maturity is regarded as a weighted average of the instantaneously short-term interest rate that is sensitive to the current monetary policy action, and an infinitely long-term rate that fully reflects the expected future path of the inflation rate. The weights depend on the speed of adjustment of the system to the steady-state inflation rate expectation and on the length of maturity of underlying instrument. The results conclude that a greater predicting accuracy is obtained from the term structure extrapolated to infer future inflation path than from the survey data.

Yield Curve as a Predictor of Future Real Activity

A number of studies have examined the link between movements in long- and short-term interest rates and macroeconomic fluctuations in real output. The evidence on the yield spread's predictive power for growth in the real GNP has been documented by Stock and Watson (1989), Chen (1991), and Estrella and Harvouvelis (1991).

The information content on term structure in predicting real economic activity may rely on a common factor explanation, in which the slope of the yield curve and future real activity are influenced by current monetary policy actions. A monetary contraction would tend to flatten the yield curve and induce a decline in real activity.

A common version of credit market theory regards the increase in yield spread as a precedence of the upswing of the real economy and inflation. Interest rates reflect the equilibrium between supply and demand conditions in credit markets. The rise in long-term yields caused by an increased demand for credit indicates stronger real growth, as credit financing facilitates increased investment and consumption expenditure.

Fisher (1907) stated that the one-period equilibrium rate of interest reveals the relative marginal value of income of today to its marginal value in next period. If a future economic downturn is expected, economic agents will interchange consumption today for a bond purchase that pays off in future periods of economic slowdown, which leads to a higher bond price and lower yield. Kessel (1956) and Fama (1986) supported the notion that the term structure moves with the business cycle. It is documented that an upwardly sloping yield curve foreshadows as well as coexists with a strong economy.

The cyclicality of the term structure is instantaneously inferred from the consumption-based asset pricing model, in which the cyclical movements in consumption produce cyclicality in term yields. In a more elaborate but expressive formal model based on the consumption capital asset pricing theory, Harvey (1988) described the information content of expected real term structure in forecasting consumption growth. In addition, the model is consistent with the observed predictability of consumption growth.

Hu (1993) formalized the link between the yield curve and real economic activity. Hu set out a model suggested by consumption-based theory and derived a closed-form solution of the term structure of interest rates. It is shown that the term spread has a linear relation with the expected growth in real output. The intertemporal equilibrium framework of Hu is shown as follows:

The model

Consider a representative consumer with a logarithmic preference receiving a single physical commodity in an infinite-horizon economy with a single production technology. The consumer can choose to consume this commodity or invest the assets.

The agent's expected lifetime utility is:

$$U(C)=E\{\int_0^\infty e^{-\rho t}U[C(t)]dt\} \tag{3.3}$$

where C(t) is the rate of consumption at time t, E(.) is an expectations operator, ρ is the consumer's constant time discount factor, and U(.) is the logarithmic utility function.

The capital productivity disturbance is assumed to be specified by a statistic or state variable x(t):

$$dX(t)=\mu_x(X,t)dt+\sigma_x(X,t)dB_t \tag{3.4}$$

where B_t denotes the standard Brownian motion, a martingale process subject to the filtered probability space.

The aggregate output is expressed as :

$$Y(t)=Y_0+\int_0^t\mu_y(Y,X,s)ds+\int_0^t\sigma_y(Y,X,s)dB_s \tag{3.5}$$

Furthermore, the condition in which the production technology possesses stochastic constant returns to scale is presumed. The dynamics of the price of default-free zero coupon bond is described by:

$$dP=\mu_p dt+\sigma_p dB \tag{3.6}$$

The consumer's wealth portfolio comprises physical investments in production, default-free security, and consumption goods borrowing or lending at risk-free interest rate γ.

$$dW=[aW(\mu_y-\gamma)+bW(\mu_p-\gamma)+\gamma W-C]dt+aW\sigma_y dB+bW\sigma_p dB \tag{3.7}$$

In equilibrium, the consumer's expected lifetime utility (3.3) is maximized based on the intertemporal budget constraint (3.7), and the bond and loan market clear.

The market-clearing stochastic processes (P, γ, a, C) involve the following closed-form formula for the equilibrium interest rate:

$$\gamma=\mu_y-\sigma_y^2 \tag{3.8}$$

Equation (3.8) indicates the term spread incorporated market's expectations of economic fluctuations. The assumptions of logarithmic preference and Ito's production process lead to the series of the risk-free interest rate fully subject to the first two moments of the production technology. The equilibrium term structure is positively related to the expected rate of growth in real output and negatively to the risk in aggregate production.

Kydland and Prescott (1988) constructed a real business cycle model that generates a positive correlation of the real rate of interest and real output. The model is a modified neoclassical growth model designed to explain the cyclical variance of a set of economic time series and the covariance between real output and other series. The model construction takes into account the equilibrium response to technological shocks for the aggregate fluctuations. An expected positive future productivity disturbance is expected to increase future output, associated expected growth in future consumption and increases in real rate of interest, as economic agents substitute current for future consumption.

King and Watson (1995) presented a real business cycle model with endogenous money and show that the nominal interest rate is positively correlated with past and current output. Further, the model specification implies that the nominal interest rate is an inverted leading indicator when a productivity process possesses a considerable temporary component. Current increases in output associated with the expectation of inverse changes in future periods will induce a rise in expected inflation. In this case, an increase in the nominal rate will precede a slowdown in real activity.

Rational Expectations Hypothesis of Term Structure and Monetary Policy

The rational expectations hypothesis of the term structure states that long-term rates are an average of current short-term rates and expected future short-term rates. Hence, it is implied that the term spreads predict future changes in short-term rates.

Various empirical studies have examined this proposition and have found that parts of the yield spreads are useful in forecasting interest rates while other parts are not. Some of the recent research emphasize the potential role of monetary policy in interpreting the performance of the yield curve.

The Federal funds rate is regarded as the policy instrument of the Federal Reserve. The formation of the market's expectations of future funds rate depends on the reaction of market participants to Federal Reserve actions, signaling changes in the funds rate target to forecast the values of macroeconomic indicators that might influence policy actions.

Mankiw and Miron (1986) considered the behavior of the Fed essential to reason out the term structure evidence in the context of rational expectations. The Fed exerts a dominant influence on the process generating short-term market interest rates. At a certain period of time given the information affecting policy decisions, the Federal funds-rate targeting procedure by which the Federal Reserve smooths short-term rates, induces an interest rate process in which there is negligible predictable variation in future short-term rates to be incorporated into term spreads. The lack of forecasting power of the spreads from three to twelve months reflects the way in which the monetary instrument is conducted, and is not a rejection of the rational expectations theory of the term structure. Cook and Hahn (1989) found the changes in the Treasury bill rates at maturity of three-, six-, and twelve-month, in response to incidents that alter the market's funds rate expectations, to be generally consistent with the Mankiw-Miron hypothesis.

A more general monetary policy explanation for the test results is provided by Cook and Hahn (1990). The variety of funds rate expectations is generated by the reaction of market participants to monetary policy or economic news announcements which signals possible upcoming change in the funds rate target. There is normally a time gap between the release of information announcements and an actual infrequent policy change triggered by the accruing weight of new information on economic growth and inflation. If the markets' expectations of sequential funds rate over a relatively future short-term period are influenced by the information announcements, and these expectations are built into the yield curve, the yield spreads for bill rate up to three months will fluctuate to a greater extent in reactions to variations in expectations than will the spreads from three to twelve months. In other words, if the funds target is expected to change in the near future and then to persist at its new level, all prevailing rates at longer horizon will embody approximately to the same extent the expected near-term changes, leading to the spreads between these rate mainly unaltered. In these circumstances, the way in which the instrument is employed, is to a certain degree responsible for the seeming failure of rational expectations hypothesis of the term structure.

McCallum (1994) showed that, when the Fed reacts to the process of long-term interest rates, the regression estimates in studies of the rational expectations hypothesis of the term structure will be theoretically related to the Fed's policy rule. Recently, Rudebusch (1995) estimated a model that explains the varying forecasting power of term spreads in the context of rational expectations, and elucidated the link between the Fed policy and the term structure. The study of Rudebusch provides empirical support for the Cook and Hahn (1990) theory. In addition it illustrates that a modified model, in which the probabilities of positive and negative target change are different at each stage in time, will generate foreseeable variations in the target rate over the course of several months and induce considerable forecasting ability in the six-month and three-month spread.

To better illustrate the potentially important element Fed behavior might play in explaining the empirical results of tests of the rational expectations hypothesis, the interest rate targeting model of Rudebusch (1995) is presented as follows:

Baseline data-generating process:

1. The deviations of funds rate(R_t) from target($\overline{R}_t$) are determined by

$$R_t = \overline{R}_t + u_t, u_t = a + b * u_{t-1} + e_t, \tag{3.9}$$

where e_t ~ iid random variable.

2. The size of target changes evolves in accordance with

$$\begin{aligned} \overline{R}_t = \overline{R_{t-1}} + \delta_t, \delta_t &= \eta \quad \text{with probability } P_t^+ \\ &= 0 \quad \text{with probability } 1 - P_t^+ - P_t^- \\ &= -\eta \text{ with probability } P_t^- \end{aligned} \tag{3.10}$$

3. The timing of target change is presented in probabilities as

$$\begin{aligned} P_t^+ &= P^d(\tau) = P^{-+}(\tau) \quad \textit{if } \delta_{t-\tau} < 0, \text{ and } \quad P_t^- = P^s(\tau) = P^{--}(\tau) \textit{ if } \delta_{t-\tau} < 0, \\ &= P^s(\tau) = P^{++}(\tau) \quad \textit{if } \delta_{t-\tau} > 0, \qquad\qquad = P^d(\tau) = P^{+-}(\tau) \textit{ if } \delta_{t-\tau} > 0, \end{aligned} \tag{3.11}$$

where $t-\tau$ is the time period since the previous non-zero target change at τ,

$P^{--}(\tau)$ and $P^{++}(\tau)$ are the probabilities of successive change in the same direction,

$P^{-+}(\tau)$ and $P^{+-}(\tau)$ are the probabilities of successive change in the different direction.

4. Long rates are determined as maintained by the pure rational expectations hypothesis of the term structure:

$$R(n)_t = 1/n[R_t + E_t \sum_{i=1}^{n-1} R_{t-i}], \tag{3.12}$$

The construction of the model permits transitory deviations of the spot rate from the target such that the intervention of the Fed to enforce the target takes place only few times over the course of a few days. The model specification also allows for the various characteristics of the behavior of the target rate.

The changes in the daily target rate as δ_t denotes the amount of target increases or decreases conducted in steps. The relative sizes of the probabilities for positive and negative changes in the target rate determine the degree of asymmetry and the timing of target changes. The dependence intensity between the signs of successive nonzero target changes as measured by P_t^+ and P_t^-, describes the likeliness of continuous movements of many steps in the same direction or a change of a given sign followed by an alteration of the reverse sign. The durations between changes of same sign as well as of different signs vary over time and are contingent on τ, the length of time apart from the previous target change.

The nature of gradual increases or decreases is specified by the differential probabilities between positive and negative target changes over short horizons. These differences generate the smoothing process of short-term interest rates, as targets are adjusted to a limited extent in moderate, deliberate steps and promptly reversed target changes seldom arise. As a result, movements in rates at short horizons are foreseeable, leading to significant discrepancies between current and forward short rates and consequently sizable term spreads. On the other hand, if positive and negative target movements are symmetrically possible, the target rate will display a near-term random walk process, such that there will be no short-term smoothing and thus no considerable difference between current and forward short-term rates.

Similarly, target persistence is defined by the difference between the probabilities of positive and negative target changes at medium-term horizons. If the Fed maintains the target rate at a level so that a target increase or decrease is symmetrically probable (specifically, $P^{s}(\tau)=P^{d}(\tau)$), essentially there will be no foreseeable changes in the target rate over a medium-term period and thus the yield spreads between rates of mediate maturity will be negligible.

The analysis presented above suggests potential differential dynamic adjustments of yield spreads in the information content of the interest rate term structure. The current study adopted an empirical, model-based approach to illustrate that the manner in which the Fed controls interest rates affects the prediction of the term structure for future short-term rate changes and the validity of the test of the rational expectations hypothesis.

The asymmetric adjustment process in term spread was further investigated by Enders and Granger (1998), and Enders and Siklos (1998). Their test results reconfirm the expectations hypothesis of the term structure and suggest a momentum threshold autoregressive adjustment in the manner in which the movement of the yield spreads toward the equilibrium relationship is postulated as asymmetric.

CHAPTER 4. ASYMMETRY IN INTEREST RATE TERM STRUCTURE

Recently empirical studies and the observed behavior of interest rates have implied generally that the dynamic adjustments of yield spreads might be different in the presence of positive deviations versus negative deviations from the long run equilibrium. This asymmetry may arise from the differential effects of monetary policy actions on market expectations of future interest rate movements. The differences in frequencies and speeds at which term spreads are altered may also stem from biased policy actions toward additional tightening or easing, to achieve the fundamental goals of monetary policy, such as wage and price inflation, real output, employment, and exchange rate. In addition, the behavior of term structure may incorporate the inherently asymmetric process found in inflation rates and the business cycle through the comovement of nominal interest rates and various macroeconomic variables.

Urich and Wachtel (1981), and Roley (1983) provided evidence that the market reacts to money growth only when the rate of the growth is outside the target ranges of the Fed. Cornell (1982) showed that the market responds to the money supply figure more substantially when the Fed concentrates on monetary aggregates. Guirquis (1994) reported that the ex-ante real and nominal interest rates respond asymmetrically to positive and negative money innovations. Enders and Siklos (1998), and Enders and Granger (1998) demonstrated this asymmetric nature of error correction among interest rates of different maturities with the data of U.S. yields and Eurorates. In the following sections, explanations are provided for asymmetry in term structure which are consistent with the rational expectations hypothesis.

Monetary Policy and the Asymmetric Dynamic Process in Term Structure

The market perception of how the monetary authority intends to respond to shocks induces asymmetries in term structure. The policy anticipation hypothesis states that, when a tight monetary policy is perceived by the markets as a central bank reaction to balance off unexpected positive shocks in money stock, the slope of the yield curve will become flatter. On the other hand, the expected inflation hypothesis states that, when the markets expect the central bank will react to accommodate unexpected positive money shocks, the slope of the yield curve will become steeper.

The fluctuations in long-term interest rates mainly reflect shifts in the current target rate of the monetary authority that anchors short maturity rates and the shifts in expectations of inflation. A tightening of policy potentially shifts both components of the long-term rate by inducing rises in short-term rates and changes in the expectation of a long-run period of inflation. Short-term rates can be directly and considerably influenced by policy actions, because open market operations and discount window lending by the central bank affect the aggregate supply of bank reserves. The reaction of long-term rates to policy actions, nevertheless, depends on the alternation of market expectations on the future path of the monetary policy. If market participants have confidence in the central bank's disinflation, long-term rates will not rise, and thus yield spreads will become narrow. However, if the disinflation is not credible, a delayed policy move might trigger a crisis of confidence that stimulates a higher inflation, and increases in short-term rates under the tightening would also employ an instantaneous and momentous upward pull on long-term rates, leading to a rise in the spread. Accordingly, yield spreads at different maturities could display an asymmetric response to the central bank's disinflation actions, contingent on the state of the central bank.

The monetary authority sometimes adopts an aggressive strategy to promote real growth or reduce a proceeding trend rate of inflation. Under this strategy, the net impact of policy actions on long-term rates is complex. The real rate effect shifts the long-term rates in the same direction as the interbank reserve rate which serves as the policy instrument, while the inflation effect shifts the long-term rates in the opposite direction.

Consider an aggressive increase in the interbank reserve rate undertaken to curtail the trend rate of inflation. When the variability of the response of long-term rates to policy actions is systematically related to the business cycle, an interbank rate tightening could have differential effects on long-term rates and, consequently, on spreads at different horizons over the business cycle. A policy tightening employs two conflicting impacts on long-term rates early in a business cycle, because financial market participants view the current actions to be followed by a sequence of consistent policy moves when an economic upswing and higher inflation are foreseen. The cyclical interbank rate increase exerts an immediate and near-term pull on long-term rates, but it may also reduce the expected rise in inflation as well. For an economy downturn that is relatively short with policy tightenings that are a little excessive, the interbank rate effect may prevail over the inflation effect. Long-term rates would probably react to a tightening to the same or even a greater extent than short-term rates, leading to increases in the yield spread during the ensuring upturn. On the other hand, the current policy stance of a tightening late in business cycle may be considered to be short-lived and likely to be reversed since an economic showdown and lower inflation are anticipated. Under these circumstances, long-term rates may present limited response or even a decline, leading to a narrowing of the yield spreads close to cyclical troughs.

Correspondingly, an aggressive reduction in the interbank rate intended to resist a recession may be particularly effective in pulling down long-term rates in the early stages of a cyclical downturn. Later, near the cyclical troughs, however, market participants may foresee an economic recovery and higher inflation and, consequently, they may regard any further persistent easing as being likely to be reversed if the economy strengthens. As excessive easing proceeds, a rise in expected inflation may offset, to a progressive extent, the immediate effect of a cyclical interbank rate on long-term rates, or even reverse the direction and cause the rise of long-term rates. Hence, long-term rates may move in a different direction from the interbank rate and short-term rates for periods near recession troughs.

The asymmetry of yield spreads may also arise from an asymmetric policy response of the monetary authority to various economic indicators, such as inflationary pressures, unemployment level, or exchange rate movements. When a central bank's primary concern is the acquisition and maintenance of credibility for its commitment to low inflation, a tightening of policy would be given more weight than an easing of policy. If policy actions, in the short run, have considerably stronger influence on the evolution of expected future short rates, a relatively small fraction of policy changes may be transmitted to inflation expectations. In this case, the expected future path of short-term rates would be built immediately into the term structure of interest rates. According to the implications of the expectations hypothesis, yield spread is an optimal forecast of a weighted sum of future changes in short-term rates. In such circumstances, policy actions biased toward a tightening may lead to sustained short-term rate increases and, thus, asymmetric behavior in yield spreads.

Type of asymmetry

The asymmetry in a yield spread time series refers to the cyclical adjustment process in which a phase of the cycle may differ from its mirror image of the opposite phase. Yield spreads may exhibit different types of asymmetric fluctuations which lie at the source of the interaction between a long-term rate and the rate of the policy instrument.

The purpose of this section is to focus on the argument for asymmetry by illustrating two distinguishing patterns of asymmetry that may prevail separately or concurrently among different types of asymmetric behavior. Figure 4.1 depicts two types of asymmetry for a detrended time series, with the level of variable located on the horizontal axes and the time on the vertical axes.

The first panel presents a cycle that relates to relative slopes or rates of change and compares reflections across hypothetical vertical axes positioned at the peaks and troughs. The process exhibits asymmetrically steep movements in a series, with a different degree of autoregressive decay depending on the variable's rising or falling.

Steepness can be generated by a model in which a recession-fighting stance of policy is much more decisive than the anti-inflation actions. The central bank promptly and substantially reduces interest rates to resist a recession when reacting to negative economic disturbances, but raises short-term rates in steps to restrain inflation such that the consequences of policy tightenings and the need for each incremental rate change can be sequentially evaluated.

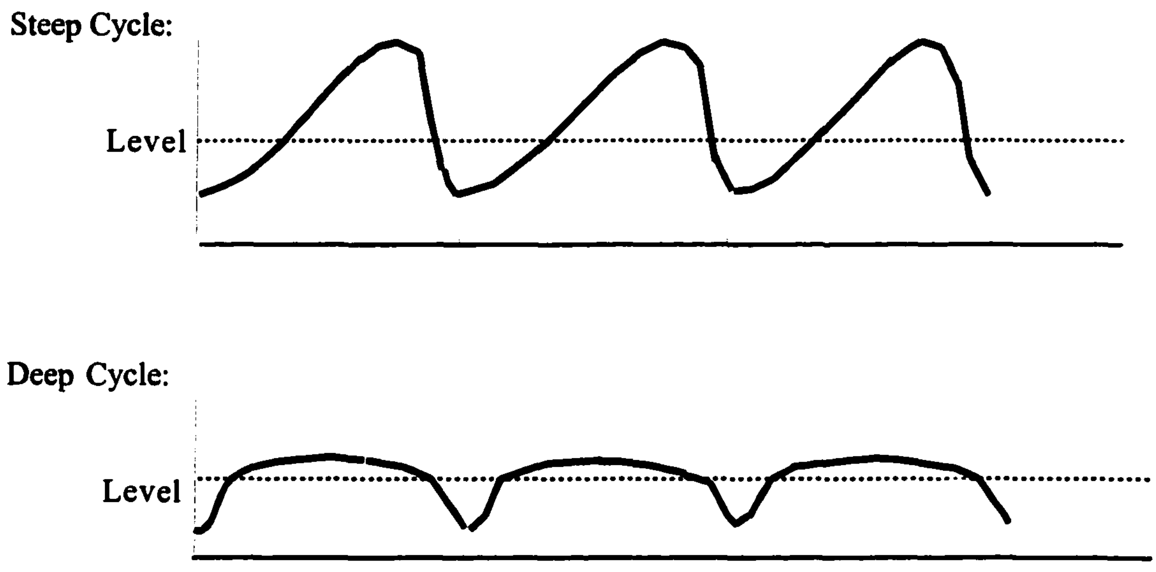

Figure 4.1. Assymetry for a detrended time series

The second panel presents a cycle which relates to relative levels and compares reflections across a hypothetical horizonal axis. The process exhibits deep movements in a series, with the speed of autoregressive adjustments depending on the variable's status above or below the trend.

Deepness can be generated by a model with an asymmetric response of long-maturity interest rates. In an economic environment with a conscientiously and credibly maintained price stability or low inflationary pressure, the upward shifts in inflation expectations in response to an interbank rate easing are likely to be more limited or more sluggish than the downward shifts in response to a tightening. A reduction in the interbank rate will have considerable effectiveness in influencing long-maturity rates and, consequently, yield spreads, since the immediate and near-term effect of the interbank rate would likely dominate the effect of inflation under these circumstances.

In contrast, an increase in the interbank rate will have a relatively small or lagged impact on long-maturity rates and yield spreads, since the instantaneous response of long rates to the interbank rates may, to a great extent, be offset by the fall in expected inflation. Moreover, the preemptive strategy of the central bank may naturally be to adopt aggressive easings in a rather short period to effectively reduce interest rates and encourage real growth while holding the line on moderate inflation expectations and, alternatively, to undertake cumulative tightenings in a sequence of actions to retain credibility for a restraint to inflation while cautiously keeping in check any real economic activity.

Inflation Persistence

An alternative interpretation of asymmetry in term structure is provided by Enders and Siklos (1998). Empirical evidence of disinflation often suggests that high inflation rates persist for several years during the first stage of disinflation, followed by a relatively rapid fall in the last few years. Since the movements in nominal interest rates at longer-term horizons reflect fluctuations in expected inflation, the comovement of nominal interest rates and expected inflation indicates the inherently asymmetric process of inflation rates which are depicted by the asymmetry found in the term structure of interest rates.

The Fisher (1930) equation states that the nominal interest rate is the sum of the constant real rates and expected decline in the purchasing power of money. Along with assertion of the expectations hypothesis that the slope of yield curve contains information about market expectations of future interest rate movements, the Fisher effect implies that the slope of term structure also contains information about the expected path of inflation.

In an economy under which the real rate of interest is determined by the equilibrium of demand for and supply of lending in the absence of taxes on interest or investment income, the Fisher neutrality postulates that nominal interest rates move one-for-one with changes in expected inflation, leaving ex-ante real rates unaltered. As inflation erodes the purchasing power of money, the nominal market interest rate embraces a competent inflation premium as a compensation to lenders.

Empirical evidence often suggests a linkage between the robustness of the Fisher effect and the persistence of inflation. If inflation is persistent and chronic in nature, then variations in actual inflation are likely to be embodied in expectations of future inflation. Based on this groundwork, nominal interest rates would tend to rapidly adopt fluctuations in the expected inflation rates, resulting in common trends between nominal interest rates and inflation.

The time series properties of the inflation rate are, to a certain degree, related to the extent to which monetary authorities accommodate inflationary shocks. The intensity of monetary accommodation is attributable to institutional structure, the anti-inflation discretion of the central bank, and disinflation costs.

Inflation persists if anticipations of additional inflation are corroborated by policy actions. Alogoskoufis and Smith (1991) provided a two-country macroeconomic model, with forward-looking price setters and a staggered wage setting to illustrate that higher monetary accommodation of inflation and exchange-rate accommodation of inflation differentials among countries increase the persistence of relative inflation rates. In their analysis of the relationship between the degree of inflation persistence and the essence of the monetary and exchange-rate accommodation, Alogoskofis and Smith also constructed a two-country demand-supply aggregate model. The demand side of economies are characterized by the equations of (1) savings and investment balance conditions (IS curves), (2) money market equilibrium conditions (LM curves), and (3) an uncovered interest-parity condition, which represent a conjecture of the perfect substitutability of financial assets between countries.

The supply side of each economy is specified by the following:

Production is related to employment proportionally as

$$y_{it}=l_{it}+q_{it} \tag{4.1}$$

where y = the firm's output,

l = the employment,

q = the marginal productivity of labor input, which is exogenous by construction,

i is a firm indicator.

The demand curve for each monopolistically competitive firm slops downwards. The discrepancy in output demand between an individual firm and the average demand corresponds to the firm's price relative to the average price.

$$y_{it}=y_t-\lambda(P_{it}-P_t) \qquad \lambda > 1 \tag{4.2}$$

where λ is the price elasticity of demand,

P Is the domestic price.

The optimal pricing rule for a monopolistically competitive firm consists of a fixed markup on unit labor expenses.

$$\overline{P_{it}}=v+w_{it}-q_{it}, \qquad v=\ln(\frac{\lambda}{\lambda-1}) \tag{4.3}$$

where w is the labor wage.

Under the proposition of the convexity of adjustment costs, the firm's intertemporal loss function imposes a penalty on price inconsistency with the static optimal price in (4.3) and price variation above a certain steady-state inflation rate Π.

$$\Lambda_{it}=E_t\sum_{s=0}^{\infty}\beta^s[\frac{1}{2}(P_{it+s}-\overline{P_{it+s}})^2+\frac{\theta}{2}(P_{it+s}-P_{it+s-1}-\Pi)^2] \tag{4.4}$$

where β = the time discount factor,

θ = the weight on the relative costs of marginal price adjustment and price inconsistency with the optimal steady-state price.

Conditional on price sluggishness and information at the end of time t-1, the firm sets wages before prices are determined in an attempt to obtain a real wage target, by taking into account anticipated inflation and labor-market tightness.

$$w_t=E_{t-1}[P_{ct}-\varepsilon U_t]+\omega_t$$
$$\equiv E_{t-1}[6P_t+(1-6)(e_t+P_t^*)-\varepsilon(N_t-l_t)]+\omega_t \tag{4.5}$$

where P_c = the consumer price,

U = the unemployment rate,

ε = the sensitivity of real wage to unemployment rate,

6 = the proportion of domestic goods for consumption,

e = the exchange rate, the price of domestic currency in terms of foreign currency,

P^* = the price of foreign country,

N = the labor force,

ω = the full-employment reservation real wage, and

N and ω are exogenous by construction, and all model variables are represented in logarithm form.

Prices are sluggish but price-setting firms are anticipatory in their conduct of setting price. Wage setting is determined by expected prices and unemployment. On account of the cost of adjustments, price changes of the firm are obtained by a sequence of moderate adjustments in stead of unusual substantial movements. As a result, a shock that raises the present price level also leads to subsequent price rises in future periods. Accommodative monetary actions succeeding a price shock generate expectations of the future rate of unemployment lower than in the absence of accommodation which, in turn, lead to higher current and expected future nominal wages. Given the convexity of adjustment costs, an optimal price-setting for the firm is to reflect to some extent on higher future nominal wages in the process of price movement from today forward, causing an initial rise in prices to persist further.

In a study of central bank behavior, Ball (1991) applied a time-consistent model approach to illustrate that if only accommodative monetary policy were conducted in reaction to temporary and exogenous macroeconomic shocks that set off initial price increases, then the temporary nature of inflation would be transformed into persistent features.

A monetary accommodation of inflationary shocks is often times motivated by high disinflation costs in terms of output loss. Chadha, Masson, and Meredith (1992) investigated the linkage between disinflation costs and the credibility of the policy actions to restrain inflation. They noted that the flexibility of labor markets which reflects the responsiveness of prices and wages to demand will also moderate the costs of disinflation.

Business Cycle Asymmetry

The behavior of macroeconomic variables over periods of the business cycle has frequently been an area of research interest. The asymmetry found in business cycles, in particular, has attracted recent interest. Several studies have emphasized the asymmetric rate of change of business cycle variables over phases of increase relative to phases of decrease. The concept of asymmetric behavior in business cycles was investigated by Keynes (1936), which indicates that contractions in an economy are often more violent but swifter than the expansion. Hamilton (1989) presented a nonlinear, asymmetric model specification for real growth rates of GNP that prevails over linear models. Brock and Sayers (1988) investigated postwar industrial production and found evidence of nonlinear structure.

Asymmetric adjustments are characterized as deep versus steep movements by Sichel (1993). Evidence found in U.S. unemployment, industrial production, and GNP indicates deepness in the series. On the other hand, evidence of sharpness is found only in unemployment. Deepness relates to cyclical behavior in which troughs are found below a course more than peaks are found above. Steepness is described as the feature depicting that contractions are steeper than expansions.

The propositions mentioned in De Long and Summers (1988) link deepness to credit problems, which suggest the inherent nature of asymmetric fluctuations in the nature rate. The bank failure generates negative macroeconomic ramifications, while there is no reciprocal consequences on positive aspects. Assume that the strength of financial institutions is determined by the discounted value of previous unanticipated variations in the collateral value of the assets underwriting their portfolios, and the discount element is determined by the flexibility by which banks reconstruct their real capital and reserves from a disturbance. In these circumstances, negative discrepancy from anticipated inflation would affect economic activity to a greater extent than positive discrepance from expected inflation.

Alternatively, the theoretical arguments of Chetty and Heckman (1985), and Baldwin and Krugman (1986) relate steepness to the asymmetric costs of upward and downward adjustment. Chetty and Heckman presented a model for estimating the lag structure of output and factor demand functions. The aggregate lag structure is determined by prices and wages, since the distribution of productive units in an operation is subject to entry and exit decisions by heterogeneous plants and firms. As a result of costless shutdowns by a productive unit and a staggered optimal investment, economic fluctuations are asymmetric with sharp declines but slow expansions. Balwin and Krugman showed that an incidental, massive exchange rate shock, which causes firms either to enter or to leave, will shift industry to another segment of the import schedule that relates to the firm's behavior and industry import volume to the exchange rate. This large disturbance may, therefore, lead to an apparent structural change in the relationship between exchange rate and import. The main reason is that the entry costs are sunk, and not all of the foreign entrants induced by a past temporary rise in the exchange rate will exit the domestic market when the exchange rate restores its initial level. On this basis, a temporary overvaluation is succeeded by a sequence of persistent reduction in the equilibrium exchange rate, which enables the retaking of lost markets.

CHAPTER 5. MONETARY POLICY IMPLEMENTATION IN JAPAN

Intermediate Targets and Instruments

Since the late 1970s, the Japanese monetary structure has gradually developed from a policy mechanism based on credit control at regulated interest rates, aimed to influence the lending of financial institutions, to a monetary-based control on broadly defined money stocks at flexible interest rates under liberalized markets of financial and foreign exchange. The primary policy instrument available to the Bank of Japan (BOJ) in the past was a so-called window guidance or moral suasion, whereby the BOJ directly controlled the amount of bank loans. However, since the 1970s, the financial system began a transition from an administratively governed structure to one that allows market forces to take a greater part in the determination of credit allocations.

Certificates of deposits (CD_S) were introduced and developed, while gensaki transactions grew substantially in volume and became the largest open market in the 1970s. The deregulation of the euro-yen market since 1980, under the New Foreign Exchange and Trade Control Law, fostered interest rate arbitration between domestic and euromarkets. Financial liberalization increased market responsiveness of interest rates and portfolio settlements, and enhanced the transition process of monetary policy into money markets and monetary aggregates.

Over the course of time since mid-1978, interbank rates became a key short-term operating target for the Bank to attain a monetary aggregate design. The Bank started to rely on changes in lending at the discount window and operations in the interbank money market, to control the supply of reserves to the banking system and, consequently, the interest rates. Furthermore, during periods of monetary tightening or easing, the changes in the Bank's stance of monetary policy and its corresponding new target levels for interbank rates have been achieved by a "reserve progress ratio" strategy, by which the Bank changes the time path of reserve supplies within one reserve accounting period.

In more recent periods, as the development of financial liberalization has proceeded in domestic markets and the growth of offshore yen markets has continued in Europe and in Tokyo, this liberalization basically has rendered the demand and supply of money to become more unstable. Thus, financial liberalization in Japan has undermined the attachment of intermediate targets to the fundamental goals of monetary policy.

In the summer of 1988, the discrepancy between interbank and open-market rates enlarged, since the expectations of subsequent monetary tightening raised short-term open-market rates, such as the CD and Euroyen rates, while interbank rates remained relatively low under the operations of the Bank. Later, in November, the BOJ countered these circumstances by further liberalizing transactions in its interbank markets and the arbitrage between interbank and open markets, and by amending money market operations and intervention procedures. The money market reform was aimed at allowing the Bank to retain its controls on short-term rates in money markets, to increase the degree of arbitrage between domestic and offshore markets as well as to encourage more free determination of interbank rates. The BOJ came to designate bill discount rates at shorter horizons, specifically one to three weeks, as key operating instruments, instead of its previous rates at horizons of one to three months. The BOJ also claimed to adopt a policy design of more frequent and adjustable operation of the official discount rate.

As a result of these reforms, the discrepancies between bill rates and Euroyen or CD rates were reduced. In addition, money market rates such as the two-month bill discount rate, which became to depend mainly on the supply of and demand for reserves, came closer to its related free market rates. The money market rates were also accorded to an officially-induced level of a one to three-week bill discount rate. The BOJ appeared to progressively emphasize controls over short-term interest rates and direct impacts on the economy, instead of its former rigid monetary target implementation policy.

Policy Formulation and Implementation

The BOJ has depended on its changes in lending at the discount window and operations in the interbank and open money markets as policy instruments to affect the reserve status of the banking system. The operations of the BOJ were directed toward offsetting or complementing the effects of the shortage of surplus of funds in the money market, as estimated by the sum of vault cash, currency held by the public, and treasury deposits.

The reserves management technique used by the BOJ primarily involves the operation of a reserve progress ratio. Japan's financial institutions are required to retain deposits in non-interest bearing accounts at the central bank in a specific proportion to deposits and other liabilities. The relative amount is determined by the multiplication of the reserve ratio and average deposits outstanding in a calendar month. On the final day of each reserve maintenance period, a course of time from the sixteenth of each month to the fifteenth of the following month, each financial institution is obliged to satisfy its legal requirements calculated for the previous month. In a sense, the system of Japan defers to a composite of a lagged reserve system and a contemporaneous reserve system. The reserve progress ratio represents the accumulative amount of prevailing daily reserves since the beginning of the present reserve accounting period is relative to the required reserves of the period.

The BOJ conducts a monetary tightening or easing by changing the time path of reserve supplies within a maintenance period. The changes in the rate of increase in the reserve progress ratio force commercial institutions to become a borrower or lender of reserves in interbank market, and thus affect call and bill discount rates.

The reserve requirement system of the BOJ was established in 1957, as an instrument directly influencing the liquidity position of commercial banks. In recent years, this policy instrument has seldom been used by the BOJ to influence monetary position in the economy. Nevertheless, the system provides a fundamental structure for market operations by the BOJ.

The BOJ has long targeted the interbank rates in its daily operations. To establish a daily market equilibrium and to avoid volatile interest rate movements, the Bank extensively uses changes in direct lending at the discount window to financial institutions. Discount window lending is an essential daily instrument for the BOJ. The discount rate charged on lending by the BOJ has been set consistently at a level lower than the call and bill rates. Accordingly, discount window lending has been rationed in Japan. This lending position is modified at the initiative of the BOJ, not of commercial banks. In addition, the effective interest charged on lending is considered as a fine to member banks for loans at very short horizons, since the interest is counted over the horizons of the loan plus one day.

The BOJ intervenes in money markets to accommodate short-term fluctuations in demand for high-power money at a target level of interest rate, or to affect commercial bank reserve conditions directed towards changing the level of target. The operations of the BOJ generally involve the bills or, since 1986, certificates of deposits. The operations also include transactions in long-term government bonds, in accordance with secular changes in the demand for money accompanied by economic growth.

The official discount rate also serves as an effective policy instrument for money market operations. The announcement of a change in the official discount rate has substantial influence on market expectations of policy stance, since statements of the BOJ are mostly accompanied by an increased intervention of the Bank in money markets.

Japan's Experience in Setting Monetary Policy

Japan has experiences three phases in setting monetary policy since 1975.

1975-1985 "money-focused"

Before the mid-1970s, there was no general consent in Japan on the consequences of money supply. As in other countries, policymakers were influenced by rampant inflation associated with disruptions in the real economy in the 1970s, followed by a new monetarist economics related to the Chicago school. In July 1975, the BOJ leaned towards the determinant impacts of money supply on inflation and the use of broadly defined money stock as an intermediate policy target. The BOJ began publishing "forecasts" of money supply in 1978, which in its design proved effective in reducing the growth rate of the broad money aggregate (M_2+CD_s) around a rather foreseeable medium-term trend, leading to greater price stability under the new policy regime of this period. Even though the forecast in general tended to accommodate the actual deviation of monetary growth from the forecast prevailing in the preceding period, the policy procedure successfully curbed money-supply growth and eradicated inflation. The success of this monetary policy may be attributed to the decisive policy tightenings of the BOJ in reaction to the 1979 oil-price shock, which prevented considerable inflation and recession, and led to macroeconomic stability in Japan. This success may also be ascribed to the money forecasts which curtailed monetary uncertainty and permitted relatively more stable and foreseeable policy actions.

1985-1993: asset inflation and asset deflation

The Japanese economy of the second half of the 1980s was marked by significant growth in economic activity, domestic currency appreciation, persistent liberalization in financial markets, and modest Consumer-Price-Index inflation. This accelerated growth was accompanied by excessive increases in asset prices that began in 1986. Nevertheless, concealed from the asset bubble was a sharp deterioration in the quality of bank balance sheet stemming from the imprudent expansion of bank credit to real-estate and equity markets, in which loans were backed by little or no collateral except the anticipation of future up-movements in price.

The burst of the asset-price bubble began in May 1989, when the BOJ raised its official discount rate. This collapse of asset prices had an adverse influence on the real economy and on the financial system. Japan's economic recession deepened due to a persistent deceleration in final demands and subsequent inventory adjustments. Nonperforming loans exceeding 10 percent of the GNP disrupted the financial system and led to the most sever financial crisis in Japan since 1927.

The sharp deceleration in asset prices in the early 1990s, following the sharp acceleration from 1985 to 1990, was considered to be driven by the self-correcting mechanism inherent in stochastic speculative processes. Market expectations of future price movements generated buying or selling, which in turn led to the sharp deviations of prices from the long-term equilibrium values in the economic bubble. The rate of monetary growth increased and the real rate of the interest fell in the early stage of asset-inflation, from the third quarter of 1985 to the second quarter of 1987. From the second quarter of 1989 to the middle of 1990 when monetary policy began a tightening, the real interest rate rose quickly and substantially. The interest rate began to fall in early 1991, as the monetary policy became relatively expansionary to counter the burst of the economic bubble. Nevertheless, the growth of monetary aggregates remained on a declining trend until the end of 1992, leading to prevailing anticipations of general price deflation and further economic slowdowns.

The BOJ's monetary policy had influential impacts during the period of asset inflation and deflation. From 1986 to 1987, monetary easing induced by discount-rate cuts was excessive in frequency and magnitude, while monetary tightening due to discount rate increases in 1989 was procrastinated and late. The money supply accelerated remarkably in 1987 and 1989, deviating from a trend of slow reduction of the money supply in previous period of 1975-1986.

The far-reaching deregulation of financial markets and the introduction of new financial instruments made the interpretation of money-supply changes difficult, since the massive monetary growth might merely rose from large-scale portfolio shifts, as individuals and corporations shifted their assets among the different financial instruments seeking the best terms and positions. In addition, the broadly defined consumer price inflation remained low, mainly due to the falls in import prices associated with currency appreciation.

The BOJ argues that its acceleration in money growth is attributed to a composition of low interest rates, strong transaction demand related to prosperity in real activity and asset markets, and fast deposit growth and bank credit expansion based on increased financial liberalization. The exchange rate of the yen against the dollar became another main factor in the design of monetary policy. In its efforts to coordinate international intervention operations and to stabilize yen appreciation, the BOJ at certain times adopted an expansionary policy stance in the latter 1980s. A rapid appreciation of the yen occurred between February 1985 and August 1986, from 360 to 260 per dollar. The Plaza Agreement of September 1985 supported further dollar depreciation and announced the commitment to policy coordination for the countries in the Group of Five. The Louvre Accord of February 1987 affirmed the cooperation of the monetary authorities of the G7 countries in maintaining the exchange rates at "current" level. The Japanese monetary policy of 1986 and 1987 was designed to restrain the yen from over-appreciation that arose from these coordinated operations and, thus, normally involved substantial foreign-exchange interventions, which in turn led to rises in the liquidity of the banking system and monetary growth. As a consequence, the foreign exchange reserves of Japan during the period of 1986-1988 escalated from $20 billion to $90 billion, accompanied by an increase in monetary-base growth from 5 percent to nearly 15 percent.

The following episode of yen appreciation refers to the change of the yen from 160 per dollar in April 1990 to below 80 per dollar in April 1995. However, since 1989, a relatively restrictive policy stance resisting asset inflation remained until the middle of 1992, in spite of sharp declines in stock and land prices as well as continued yen appreciation.

1993-present

An economic slowdown in Japan became apparent by the middle of 1991. Moderate discount rate cuts were employed in 1991 and 1992, but the actions did not stimulate the economy. In February 1993, the official discount rate was reduced to the historically low level of 2.5 percent as that between 1987 and 1989. Massive and consistent monetary actions started around April 1993. Nevertheless, the economic outlook for a recovery was soon dampened by further appreciation of the yen in the first half of 1993, from 125 to under 110 per dollar. The growth of the monetary base and the growth of M_2+CD_s continued to decline sharply, as recession and the collapse of assets prices followed. Japan's fiscal policy became expansionary in 1994 and 1995, owing to budget revisions in the middle of each of these years.

Intervention operations in the foreign exchange market were adopted by the BOJ to resist the rise of the yen, and were particularly supported by the coordinated policy actions of other monetary authorities in 1994, after the April meeting of the G7 countries' central banks and finance ministers. However, the continued appreciation of the yen in 1994 and 1995 resulted in the deterioration of domestic investment and monetary growth. The BOJ countered the yen appreciation and the stagnant economy by lowering the official discount rate again to 1.0 percent in April 1995, and then to a record low of 0.5 percent in September 1995. Nonetheless, it was not clear that the stimulative monetary policies were effective, as the deflationary impacts of the collapse of the asset-price bubble and the worsening banking problem continued to hinder recovery of the economy.

CHAPTER 6. METHODOLOGY

The purpose of the study was to examine the evolving relationship between interest rates of different maturities in Japan, allowing for possible asymmetrical movements toward a long-run equilibrium rather than use of the traditional analysis of symmetrical movements. Proposed are two types of asymmetric test methodologies in the form of threshold autoregressive and momentum threshold autoregressive adjustment representations to account for the time-series features of deepness and steepness of slope in interest rates. In this chapter conventional unit-root and cointegration tests are introduced first, followed by asymmetric test methodologies.

Conventional Unit-Root and Cointegration Tests

Dickey-Fuller test

The procedure to determine whether a time series follows a random walk or stationary process is based on the unit-root hypothesis test. Standard unit-root tests assume linearity and symmetric adjustments. Consider a series generated from the following first-order process:

$$S_t = \rho^* S_{t-1} + \varepsilon_t \; ; \qquad \varepsilon_t \sim \text{IID}\,(0, \sigma_\varepsilon^2) \tag{6.1}$$

If $\rho^* = 1$, the $\{S_t\}$ sequence follows a random walk with time-dependent mean and variance. If $-1 < \rho^* < 1$, $\{S_t\}$ is stationary. In the presence of a unit root $\rho^* = 1$, the estimate of ρ^* in the form of (6.1) will be biased below unity and thus classical statistical tests on the coefficient ρ^* will be inappropriate. Accordingly, Dickey and Fuller (1979, 1981) formulated the procedure to test for the presence of a unit root as:

$$\Delta S_t = \rho S_{t-1} + \varepsilon_t \quad ; \qquad \varepsilon_t \sim \text{IID}\,(0, \sigma_\varepsilon^2) \tag{6.2}$$

where $\rho=\rho^*-1$. Under the null hypothesis of a unit root Ho: $\rho=0$, a regression of an I(0) variable on a I(1) variable as specified in equation (6.2) will induce a t-ratio on the coefficient estimate that does not have a standard Student-t distribution. Using a Monte Carlo simulation, Dickey and Fuller derived the distribution of the appropriate test statistics for testing the hypothesis that ρ equals zero. A few of the critical values are represented in Table 6.1.

Table 6.1. Empirical cumulative distribution of τ for the Dickey-Fuller unit root test

	Probability of a Smaller Value							
Sample Size	*0.01*	*0.025*	*0.05*	*0.10*	*0.90*	*0.95*	*0.975*	*0.99*
250	-3.46	-3.14	-2.88	-2.57	-0.42	-0.06	0.24	0.62
500	-3.44	-3.13	-2.87	-2.57	-0.43	-0.07	0.24	0.61

Constant τ_μ

The sequence $\{S_t\}$ refers to the yield spread series for the hypothesis testing employed in this study. Rejection of the null hypothesis of a unit root in the series $\{S_t\}$: $\rho=0$ in favor of the alternative of stationarity: $-2<\rho<0$ implies that the interest rate differential is stationary and the expectations hypothesis finds support in the empirical results.

The Dickey-Fuller equation can also be modified by (1) the use of deterministic regressors in the estimations to allow for the possibility that "nuisance" parameter enter into the series, (2) the use of an alternative Phillips-Perron test to account for the heterogeneity or weakly dependence in the error term of a unit root process, (3) the incorporation of the structural breaks in the test, and (4) the addition of the lag differences of S_t to purge possible serial correlation in the error term.

Engle-Granger test and Johansen method

An n-dimensional nonstationary series X_t which can be transformed to yield a stationary invertible ARMA representation after differencing d times is said to be integrated of order d, denoted by $X_t \sim I(d)$. Given that all components of the vector series $X_t=(x_{1t},x_{2t},\ldots,x_{nt})'$ are integrated of order d, there may exist linearly independent vector $\beta_i=(\beta_1^i,\beta_2^i,\ldots,\beta_n^i)'$, $i = 1,\ldots,r$, such that linear combination $\beta_i'X_t= \beta_1^i x_{1t}+\beta_2^i x_{2t}+\ldots+\beta_n^i x_{nt}$ is integrated of order (d-b), b>0, denoted by $X_t \sim CI\ (d,b)$, with cointegrating rank r. The number of linearly independent cointegrating vectors β_i's that span the cointegrating space can be at most n-1 for n-dimensional series X_t.

Conventional cointegration model assume symmetry and linear relationship. For a system of time series of which each are individually integrated of order 1, the Engle and Granger (1987) methodology of cointegration test involve two procedures. The first step is to estimate the long run equilibrium relationship among the variables of interest in the form of :

$$R_{1t} = \beta_2 R_{2t} + \beta_3 R_{3t} + \ldots + \beta_n R_{nt} + \mu_t \tag{6.3}$$

where $R_t = (R_{1t}, \ldots, R_{nt})'$ and all of the components of R_t are integrated of order 1. μ_t is the stochastic disturbance term. In the analysis of interest rate term structure conducted in this paper, R_t refers to a 2-variate time series of the yields of different maturities, and there can be at most 1 linear cointegration vector.

The second step follows the Dickey-Fuller type test for the presence of a unit root to exam the order of the integration of the residuals. If the series are not cointegrated, then there must be a unit root in the residuals. If the series are cointegrated, then the residuals will be stationary. The test is conducted by the estimation of ordinary least square regression of the form:

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t \quad ; \qquad \varepsilon_t \sim \mathrm{IID}\,(0, \sigma_\varepsilon^2) \tag{6.4}$$

where the estimated residuals from the cointegrating regression of equation (6.3) are used in equation (6.4). The null hypothesis of non-cointegration is Ho: $\rho = 0$. If the error term of a unit root process of equation (6.4) does not appear to be white noise, an augmented Dickey-Fuller test of the regression residuals, which includes the lagged changes in the $\{\mu_t\}$ sequence, can be used instead of (6.4).

The stationarity of the residuals requires $-2 < \rho < 0$. The rejection of the null hypothesis of a unit root in the residual series implies that $R_{1t}, R_{2t}, \ldots, R_{nt}$ are cointegrated in a system with a symmetric adjustment mechanism toward the attractor-- the long run equilibrium $\mu_t = 0$, and the support for the expectations hypothesis is obtained in empirical results.

Engle and Granger (1987) point out that the t-ratio from this test will not have the Dickey-Fuller distribution when the true parameter β_i is unknown. The OLS estimation in equation (6.3) will choose $\hat{\beta}_i$ which minimizes the residual variance and yield a estimated residual series more stationary than the true ones. The use of the test statistics of the Dickey-Fuller distribution will reject the null too often. Therefore, alternative distributions of the test are proposed and tabulated according to Engle and Granger (1987), Engle and Yoo (1987, 1990), and Mackinnon (1990). The Engle and Yoo (1987) critical values for 2-variable cointegration test with a sample size of 200 are depicted in Table 6.2.

Table 6.2. The Engle-Yoo critical values for the cointegration test

		Significance Level		
No. of Variables	Sample Size	*1%*	*5%*	*10%*
2	200	4.00	3.37	3.02

Alternatively, the Johansen (1996) and Stock and Watson (1988) methodology for the cointegration tests can be expressed in a form as:

$$\Delta R_t = \pi R_{t-1} + \varepsilon_t \tag{6.5}$$

where $R_t = (R_{1t}, ..., R_{nt})'$ $n\times 1$, and all of the components of R_t are integrated of order 1. π is a matrix of $n\times n$ dimension. $\varepsilon_t = (\varepsilon_{1t},...,\varepsilon_{nt})'$ $n\times 1$, and ε_t's are normally distributed but may be correlated simultaneously.

The null hypothesis of non-cointegration is that the rank of π equals 0. The test procedure involves the estimation of π and the determination of the rank of π. As with the Dickey-Fuller test, equation (6.5) can also be generalized to allow for the presence of a drift term and a higher-order autoregressive process.

Error correction representation

In vector autogregression analysis (VAR), the set of variables is treated symmetrically. The feedback by which each variable responds to the effects from others are incorporated in the VAR model to account for the dynamic adjustment. If a nonstationarity of the series exists but is not accounted for in a VAR model, spurious regression estimates with no economic meaning will be induced. Conventional differencing which involves calculating successive changes in the values of a data series is commonly used to transfer nonstationary series into stationary ones. However, according to Engle and Granger (1987), a model which differences the relationship while the nonstationary time series are cointegrated may entail a misspecification error. A temporary shock to the cointegrated series will be corrected through the subsequent periods, so that the series will not drift far away from its long-term equilibrium relation. Granger and Engle proved that cointegrated series have a corresponding error correction representation and that error correction models generate cointegrated series. Johansen (1988) also showed that, at the presence of nonstationarity and cointegration of the series, the level VAR model can be reparameterized as the following Error-Corrected VAR model to include the deviations from the equilibrium as an explanatory power of the movement of the series.

$$\Delta R_t = \theta D_t + \Gamma_1 \Delta R_{t-1} + \ldots + \Gamma_p \Delta R_{t-p+1} + \Pi R_{t-1} + \varepsilon_t , \tag{6.6}$$

where Δ is the difference operator

$\varepsilon_t = (\varepsilon_{1t}, \ldots, \varepsilon_{nt})$ $n \times 1$, $\varepsilon_t \sim N(0, \Omega)$

D is the deterministic mean and seasonal constants

θ is $n \times 1$, Γ_i's are $n \times n$, and Π is $n \times n$ unknown parameters.

The short-term dynamics of the VAR model are captured by matrices Γ_1 through Γ_p, while the long-term cointegration relationships among series are accounted for in matrix Π. The rank of Π is denoted as r. If the rank of Π is 0, series are nonstationary in level but stationary in first difference. The model (6.6) collapses to the VAR model by first differencing. If the rank of Π is full, the series is stationary and the level specification is appropriate. If the rank of Π is between 0 and n, the series is nonstationary and cointegrated, with the number of cointegration relationships among the series equals to r. In this case, model (6.6) is called the Error-Correction VAR model, in which there are r independent restrictions imposed on the long-term solutions. Π has a full rank factorization $\alpha\beta'$, α and β are matrices of dimension $n\times r$, and β are the cointegrating vectors. $\beta'\ R_{t-1}$ represents stationary error terms due to the short-term deviation from the long-run equilibrium cointegration relationships. α represents the speed of adjustment of series toward the cointegration relationships.

Asymmetric Regimes

Conventional unit-root and cointegration tests presuppose symmetric adjustment processes. Note that in equation (6.2), ΔS_t is strictly proportional to the absolute value of S_t, and the speed of adjustment ρ^* in the attractor $\rho^* S_{t-1}$ is regarded as invariant to S_{t-1} or ΔS_{t-1}. Similar implications can also be derived from Engle-Granger as well as Johasen tests.

The implicit assumption of symmetric adjustment is problematic since recent empirical studies and the observed behavior of interest rates have implied generally that the dynamic adjustments of yield spreads might be different from the long run equilibrium in the presence of positive versus negative deviations. This asymmetry may arise from the differential effects of monetary policy actions on market expectations of future interest rate movements. The differences in frequencies and speeds at which yield spreads are altered may also stem from biased policy actions toward additional tightening or easing, to achieve the fundamental goals of monetary policy. In addition, the behavior of term structure may incorporate the inherently asymmetric process found in inflation rates and the business cycle through the co-movement of nominal interest rates and various macroeconomic variables.

Sickel (1993) discussed two types of asymmetric cycles in time series: deepness and steepness. Deepness occurs when troughs are more pronounced that peaks and steepness occurs when contractions are steeper than expansion. In the presence of an asymmetric adjustment process, the standard unit-root and cointegration tests and their corresponding error correction representation may entail a misspecification error. Corresponding to those two variations in the adjustment processes, two specific asymmetric test methodologies in the frameworks of Enders and Granger (1998) and Enders and Siklos (1998) are introduced in the form of threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) adjustments.

Asymmetric Unit-Root and Cointegration Tests

In the presence of an asymmetric adjustment process, the standard unit-root and cointegration tests and their corresponding error correction representation may entail a misspecification error. Two types of asymmetric test methodologies in the form of threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) adjustments representations were advanced by Enders and Granger (1998) and Enders and Siklos (1998).

Enders-Granger unit-root test

The TAR representation for testing the unit-root, in which a data generating mechanism of dynamic asymmetric processes is incorporated, is represented as follows:

$$\Delta S_t = I_t \rho_1 S_{t-1} + (1 - I_t) \rho_2 S_{t-1} + \varepsilon_t \tag{6.7}$$

$$I_t = \begin{cases} 1 & \text{if } S_{t-1} \geq 0 \\ 0 & \text{if } S_{t-1} < 0 \end{cases} \tag{6.8}$$

where I_t is the Heaviside indicator function.

The M-TAR representation revises the Heaviside indicator function in equation (6.8) as:

$$I_t = \begin{cases} 1 & \text{if } \Delta S_{t-1} \geq 0 \\ 0 & \text{if } \Delta S_{t-1} < 0 \end{cases} \tag{6.9}$$

where I_t is the Heaviside indicator function.

Denote by Φ and Φ^* F-statistics for testing the null hypothesis of $\rho_1 = \rho_2 = 0$ under the TAR and the M-TAR representation, respectively. The distributions of Φ and Φ^* are determined by the form of the attractor. Enders and Granger (1998) deduced the critical values and the power tests of the Φ- and Φ^* -statistic in comparison with the traditional Dickey-Fuller test via comprehensive Monte Carlo experiment. The test models can be modified to incorporate a constant term or a drift plus linear time trend as the attractor, and to include high order augmented form of equation (6.7). The Φ- and Φ^*- statistics are tabulated in Table 6.3, while the critical values Φ_μ -and Φ^*_μ - statistics corresponding to the test of the null of a unit root with a non-zero sample mean are in Table 6.4.

Table 6.3. Empirical distribution of the Enders-Granger statistic of Φ and Φ^*

	Φ - Statistic			Φ^* - Statistic		
Sample No.	90%	95%	99%	90%	95%	99%
100	3.18	3.95	5.69	2.83	3.60	5.38
250	3.10	3.82	5.53	2.68	3.41	5.10

Table 6.4. Empirical distribution of the Enders-Granger statistic of Φ_μ and Φ^*_μ

	Φ - Statistic			Φ^* - Statistic		
Sample #	90%	95%	99%	90%	95%	99%
100	3.79	4.64	6.57	4.11	5.02	7.01
250	3.74	4.56	6.47	4.05	4.95	6.99

The percentage of times that the null hypothesis of a unit root process is correctly rejected is reported. The Φ^* -statistic often outperforms the Dickey-Fuller test when the adjustment is asymmetric in nature. However, the Dickey-Fuller test performs better than the Φ-statistic on average, since the addition of one more coefficient to the estimation reduces the power of the Φ-statistic. Nevertheless, under the multi-threshold setting, the relative power of the Φ-statistic over the Dickey-Fuller test is enhanced by the increases of the degree of asymmetry.

Enders-Siklos cointegration test

The cointegration tests incorporate the TAR and M-TAR adjustments into the unit-root tests of the residuals of the cointegration regression. The threshold autoregressive model replaces the equation (6.4) by:

$$\Delta\mu_t = I_t\, \rho_1\, \mu_{t-1} + (1 - I_t)\, \rho_2\, \mu_{t-1} + \varepsilon_t \tag{6.10}$$

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq 0 \\ 0 & \text{if } \mu_{t-1} < 0 \end{cases} \tag{6.11}$$

where I_t is the Heaviside indicator function.

The difference in the value of ρ_1 and ρ_2 permits a state-dependent autoregressive decay. The time-series feature of "deep" cyclical processes, in which positive deviations from the cointegration equilibrium are more prolonged than negative deviations, can be captured by this model. There is an accelerated autoregressive decay when the equilibrium error lies below and a sluggish adjustment when it lies above the attractor.

The momentum-threshold autoregressive model allows the autoregressive decay to depend on the change in μ_{t-1}. The model would generate asymmetric dynamic processes in which a greater extent of momentum for adjustments is present in one direction than the other. The Heaviside indicator function under the setting becomes:

$$I_t = \begin{cases} 1 & \text{if } \Delta\mu_{t-1} \geq 0 \\ 0 & \text{if } \Delta\mu_{t-1} < 0 \end{cases} \tag{6.12}$$

The M-TAR representation could capture the feature of sharp movements as denoted in Delong and Summer (1986), and Sichel (1993). In the context of the model, little decay when μ_{t-1} is increasing but significant decay when μ_{t-1} is declining would induce adjustment paths in which the increases are prolonged while decreases revert swiftly toward the attractor.

The null hypothesis of non-cointegration is Ho: $\rho_1 = \rho_2 = 0$. A sufficient condition for the stationarity in the residual series $\{\mu_t\}$ is $-2 < (\rho_1 , \rho_2) < 0$. The distribution in this case, where the underlying dynamics are asymmetric TAR and M-TAR adjustments, is nonstandard and reported in Table 6.5 and Table 6.6, respectively. These critical values for testing whether there is a cointegration relationship among two variables in asymmetric framework was computed through a Monte-Carlo simulation by Enders and Siklos (1998).

Table 6.5. Empirical distribution of the Enders-Siklos statistic of φ for the two-variable TAR model

	Probability Distribution								
	No Lagged Change			One Lagged Change			Four Lagged Change		
No.	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*
100	5.04	6.07	8.20	4.99	5.98	8.21	4.94	5.91	8.22
500	4.87	5.80	7.89	4.89	5.84	8.04	4.91	5.82	7.95

Table 6.6. Empirical distribution of the Enders-Siklos statistic of φ^* for the two-variable M-TAR model

	Probability Distribution								
	No Lagged Change			One Lagged Change			Four Lagged Change		
No.	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*
100	5.52	6.57	9.04	5.43	6.45	8.75	5.34	6.35	8.73
500	5.29	6.34	8.54	5.35	6.33	8.61	5.32	6.30	8.53

When the series of μ_t is stationary, a standard F-statistic can be used for the test of a symmetric adjustment of $\rho_1 = \rho_2$ (Tong, 1983).

The modifications to the model may also involve an alternative non-zero linear attractor and higher-order processes:

(1) a constant drift term as the attractor[1]

$$\Delta\mu_t = I_t \rho_1 [\mu_{t-1} - a_0] + (1 - I_t) \rho_2 [\mu_{t-1} - a_0] + \varepsilon_t \qquad (6.13)$$

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq a_0 \\ 0 & \text{if } \mu_{t-1} < a_0 \end{cases} \qquad (6.14)$$

where a_0 is a constant, and $\mu_t = a_0$ is the attractor.

When the stationarity condition of $-2<(\rho_1, \rho_2)<0$ holds, the system will converge to the long-run equilibrium in which $R_{1t} = a_0 + \beta_2 R_{2t} + \beta_3 R_{3t} + ... + \beta_n R_{nt}$. At $\rho_1 = \rho_2 \neq 0$, the model reduces to a symmetric adjustment specification with an AR(1) equilibrium error in the form of $\Delta\mu_t = a_0 + \rho\, \mu_{t-1} + \varepsilon_t$.

(2) A drift and linear time trend as the attractor

$$\Delta\mu_t = I_t \rho_1 [\mu_{t-1} - a_0 - a_1(t-1)] + (1 - I_t) \rho_2 [\mu_{t-1} - a_0 - a_1(t-1)] + \varepsilon_t \qquad (6.15)$$

[1]. I follow Enders & Siklos (1998) by not including an intercept in the adjustment equation.

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq a_0 + a_1(t-1) \\ 0 & \text{if } \mu_{t-1} < a_0 + a_1(t-1) \end{cases} \qquad (6.16)$$

where a_0 and a_1 are constant parameters, and $\mu_{t-1} = a_0 + a_1(t-1)$ is the attractor.

When the stationarity condition of $-2<(\rho_1, \rho_2)<0$ is satisfied, the series of equilibrium error μ_t would be trend stationary.

(3) higher-order processes

The lag difference of μ_t can be incorporated into the estimation of equation (6.10), (6.13), and (6.15), to purge possible serial correlation in the disturbance term of a unit root process. The p-th order augmented form of equation (6.10) can be expressed as:

$$\Delta\mu_t = I_t \rho_1 \mu_{t-1} + (1 - I_t)\rho_2 \mu_{t-1} + \sum_{i=0}^{p-1} \beta_i \Delta\mu_{t-i} + \varepsilon_t \qquad (6.17)$$

The condition for the stationarity in μ_t requires that all roots of the characteristic equation of $(1 - \beta_1 B - \beta_2 B^2 - \ldots - \beta_p B^p) = 0$ lie outside the unit circle.

When denoting φ and φ^* as the F-statistics for testing the null hypothesis of $\rho_1 = \rho_2 = 0$ under the TAR and the M-TAR representation, respectively, then the distributions of φ and φ^* are determined by the following factors: (1) the number of variables in the cointegrating relationship, (2) the form of the attractor, and (3) the number of lags in the regression of a unit root process. Enders and Siklos (1998), and Enders and Dibooglu (1998) provided the power tests of the φ- and φ^*- statistic in comparison with the Engle-Granger test. The power of the Engle-Granger test generally exceeds that of the φ-statistic. The φ^*- test has at least as much power as φ- test. The power of φ^*-statistic increases relative to that of the Engle-Granger test when the degree of asymmetry increases.

Error correction representation

An error-correction representation implied by the cointegration equations in the TAR and M-TAR models can be expressed as:

$$\Delta R_t = \Pi_1 I_t \; \mu_{t-1} + \Pi_2 (I_t - 1) \; \mu_{t-1} + \varepsilon_t \tag{6.18}$$

where Π_1 and Π_2 are vectors of dimension $n \times 1$, which represent the speed of adjustment of series toward the cointegration relationship.

Consistent estimates of the threshold

Since there is no a priori justification for the coincidence of the threshold with the attractor, the linear attractor of the Heaviside indicator function as applied to the study of expectations hypothesis of interest rate term structure may refer to the following forms:

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq a_0 \\ 0 & \text{if } \mu_{t-1} < a_0 \end{cases} \qquad \text{or} \qquad I_t = \begin{cases} 1 & \text{if } \Delta\mu_{t-1} \geq a_0 \\ 0 & \text{if } \Delta\mu_{t-1} < a_0 \end{cases}$$

In general, the value of a_0 is unknown and needs to be estimated along with the values of ρ_1 and ρ_2. By demeaning the $\{\mu_t\}$ sequence, the Enders-Granger test procedure actually employs the sample mean of the sequence as the threshold estimate of a_0. Nevertheless, the sample mean is a biased threshold estimator in the presence of asymmetric adjustments. For instance, if autoregressive decay is more sluggish for positive deviations of μ_{t-1} from a_0 than for negative deviations, the sample mean estimator will be biased upwards.

A consistent estimate of the threshold a_0 can be obtained by utilizing Chan's (1993) method which searches over possible threshold values to minimize the residual sum of squares from the fitted model. The residual sequence is sorted in order such as $\mu_1^{a_0} < \mu_2^{a_0} < \ldots < \mu_T^{a_0}$ where

T designates the number of observations. The smallest and largest 15% of $\{\mu_i^{a_0}\}$ values are discarded. For each value of the remaining $\mu_j^{a_0}$, the equation in the form of (6.2.7) is estimated by setting a_0 equal to $\mu_j^{a_0}$. The estimation which yields the lowest residual sum of squares provides the appropriate estimate of the threshold.

Enders (1999), and Enders and Siklos (1999) employ Chan's methodology to a Monte Carlo study to obtain the F-statistics for the null hypothesis for $\rho_1 = \rho_2 = 0$. Using the threshold value consistent with Chan's procedure, they present the distribution of the resulting F-statistics, denoted as the $\Phi_\mu(c)$ and $\Phi_\mu^*(c)$ statistic for the unit-root test in TAR and M-TAR specification, and the $\varphi_\mu(c)$ and $\varphi_\mu^*(c)$ statistic for the cointegration test in TAR and M-TAR representation, respectively (Table 6.7, 6.8, 6.9, and 6.10). Along with the F-statistic, the two t-statistics for the null hypotheses $\rho_1 = 0$ and $\rho_2 = 0$ are also recorded as t-max and t-min for the individual t-statistic of the least negative value and of the most negative value of ρ_i, respectively. A few of the t-max statistics for cointegration tests are represented in Table 6.11 and 6.12.

Table 6.7. Distribution of the $\Phi_{\mu}(c)$- statistic for the TAR adjustment

	Probability Distribution											
	No Lagged Changes				One Lagged Change				Four Lagged Changes			
No.	*90%*	*95%*	*97.5%*	*99%*	*90%*	*95%*	*97.5%*	*99%*	*90%*	*95%*	*97.5%*	*99%*
100	5.08	6.06	6.93	8.19	5.39	6.34	7.30	8.54	5.38	6.32	7.29	8.56
250	5.11	6.03	6.88	8.04	5.26	6.12	6.99	8.14	5.36	6.29	7.15	8.35

Table 6.8. Distribution of the $\Phi^{*}_{\mu}(c)$- statistic for M-TAR adjustment

	Probability Distribution											
	No Lagged Changes				One Lagged Change				Four Lagged Changes			
No.	*90%*	*95%*	*97.5%*	*99%*	*90%*	*95%*	*97.5%*	*99%*	*90%*	*95%*	*97.5%*	*99%*
100	4.81	5.77	6.73	7.99	4.77	5.71	6.56	7.90	4.74	5.70	6.67	7.97
250	4.70	5.64	6.51	7.64	4.64	5.54	6.40	7.56	4.64	5.54	6.39	7.61

Table 6.9. Distribution of the $\varphi_{\mu}(c)$- statistic for the TAR adjustment

	Probability Distribution								
	No Lagged Change			One Lagged Change			Four Lagged Change		
No.	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*
100	5.95	6.95	9.27	6.02	7.08	9.51	6.35	7.41	9.88
250	5.93	6.93	9.15	5.92	6.93	9.18	6.44	7.56	10.18

Table 6.10. Distribution of the $\varphi^{*}_{\mu}(c)$- statistic for the M-TAR adjustment

	Probability Distribution								
	No Lagged Change			One Lagged Change			Four Lagged Change		
No.	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*
100	5.73	6.78	9.14	5.76	6.86	9.29	5.52	6.56	8.91
250	5.58	6.62	8.82	5.57	6.63	8.84	5.32	6.32	8.47

Table 6.11. Distribution of t-max* for the TAR adjustment[a] for the TAR adjustment

	Probability Distribution								
	No Lagged Change			One Lagged Change			Four Lagged Change		
No.	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*
100	-1.61	-1.85	-2.35	-1.65	-1.90	-2.39	-1.66	-1.92	-2.44
250	-1.59	-1.84	-2.31	-1.61	-1.86	-2.33	-1.52	-1.73	-2.30

[a] The F-statistics will be appropriate only under the circumstances where the point estimates for ρ_1 and ρ_2 both indicate convergence. The t-min statistics was shown to have little power (Enders & Granger, 1999) and, thus, only a few of the t-max* for the cointegration test are presented here.

Table 6.12. Distribution of t-max* for the M-TAR adjustment[a]

	Probability Distribution								
	No Lagged Change			One Lagged Change			Four Lagged Change		
No.	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*	*90%*	*95%*	*99%*
100	-1.65	-1.90	-2.37	-1.67	-1.94	-2.44	-1.65	-1.90	-2.42
250	-1.66	-1.90	-2.36	-1.67	-1.91	-2.37	-1.66	-1.90	-2.37

[a] The F-statistics will be appropriate only under the circumstances where the point estimates for ρ_1 and ρ_2 both indicate convergence. The t-min statistics was shown to have little power (Enders & Granger, 1999) and, thus, only a few of the t-max* for the cointegration test are presented here.

CHAPTER 7. TESTING THE EXPECTATIONS HYPOTHESIS OF TERM STRUCTURE OF INTEREST RATE IN AN ASYMMETRIC FRAMEWORK

A number of empirical studies which examine the proposition of the expectations hypothesis find that parts of the yield spreads are useful in predicting interest rates while other parts are not. Some recent researches emphasize on the potential role of monetary policy in interpreting the performance of the yield curve. The manner in which the monetary authority controls interest rates may induce differential dynamic adjustments of yield spreads in the information content of interest rate term structure. Besides, the behavior of term structure may incorporate the inherently asymmetric process in inflation rates and business cycle through the comovement of nominal interest rates and various macroeconomic variables. To account for the asymmetry of adjustment process, the unit-root test advanced by Enders and Granger (1998) and the cointegration test methodologies by Enders and Siklos (1998) are employed in the study.

The remainder of this chapter is organized as follows. Section 7.2 provides a detailed description of the data. In section 7.3 the procedure for unit-root and cointegration tests, which allows for threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) adjustments, is described, followed by test results and interpretation of the results.

Description of the Data

The analysis has been conducted on weekly series of Euroyen deposit rates and two Japanese onshore interest rates, the Gensaki and CD rates.

The Gensaki market, an institution for bond repurchase agreements over the course of time of one to three months, has historically been the most deregulated and important short-term money market in Japan of 1970s. The Gensaki market differs from interbank markets such as those for call money and discounted bill, by allowing business corporations, government pension funds, and nonresidents, in addition to financial institutions, to participate into the market. The depth of the Gensaki market declines as the financial liberalization of Japan proceeds. The development of alternative financing instrument in short-term money market, the removal of restrictions on banks trading in the secondary bond market, and the imposition of security transfer taxes and a rise in the volatility of long-term government bonds, all contribute to the falls in corporations' demand for the Gensaki and the changes in the inventory financing of security firms.

The domestic sales of 3-month certificates of deposit (CD_S) were introduced in Japan around the middle of 1979. The Japanese CD market expanded rapidly since the mid-1980s, and now the CD rate is more representative of Japanese money market rate than the Gensaki rate.

The data on one-, two-, and three-month Gensaki rates are obtained from the Japan Securities Dealers Association. The sample period covers the August 1980 to July 1985.

The data for the one- and three-month Japanese CD rates are from the database of the Nikkei Telecom, and the spot, one-, and three-month Euroyen deposit rates are from the Bank for International Settlements, during the period between July 1985 and October 1998. Financial deregulation has decreased the depth of the Gensaki market. The fast growth of the Japanese CD market since the mid-1980s now make the CD rate a relatively good proxy for the short-term money market rate in Japan. The analysis of the first period up to July 1985 studies the Gensaki rate process, while the second and the third time intervals from July 1985 forwards investigates the CD rate behavior. Three subsets corresponding to Japanese economic events and monetary policy regimes, which may have influenced the behavior of interest rates, have been analyzed.

The first subset, which covers the period up to July 1985, is consistent with the phase of "money-focused" policy regime. The break point is consistent with the notable change in the character of the term structure in 1985 identified by Leung, Sanders, and Unal (1991), and Campbell and Hamao (1991). Short-term rates moved choppily before late 1985, then exhibited a long period of smooth fluctuation. The period of sharp appreciation of the yen and the bubble economy with the newly liberalized financial system, corresponds to the second subset that covers the period of July 1985 until January 1993. The final subset from February 1993 onwards corresponds to the policy stance of monetary expansion and a recession associated with the asset-deflation.

Test Procedure and Interpretation of the Results

Test procedure

The asymmetric unit root and cointegration tests for TAR and M-TAR adjustments are carried out using the following procedure:

Step 1: OLS estimation of the equilibrium equation.

1. The interest rate differential is regressed on a constant, for the unit-root test of the yield spread. This test imposes the (1,-1) cointegrating vector proposition.
2. One variable is regressed on a constant and the other variable(s), for the cointegration test of the term structure.

Step 2: OLS regression is performed on the residual sequence of the equilibrium relationship estimated in the first step. The indicator function I_t will be determined by the type of asymmetry under consideration. The diagnostic checking of the regression residuals from the estimation of the equation of a unit root process and the AIC/SBC criterion are undertaken to determine the lag lengths to purge the autocorrelations among the error terms.

Step 3: The sample statistics for the null hypothesis of nonstationarity :$\rho_1=\rho_2=0$, are compared with the appropriate critical values in Table 6.3, 6.4, 6.5 or 6.6. The statistic for the Enders-Granger unit-root test depends on the presence of the various deterministic regressors in the attractor, while the statistic for the Enders-Siklos cointegration test depends on the form of the attractor, the lengths of the lag, and the number of variables in the cointegration regression. If the alternative hypothesis is accepted, the restriction of symmetric versus the alternative of asymmetric adjustment can be tested using standard F-statistics.

Step 4: If the test results indicate cointegration with asymmetric adjustment, Chan's method (1993) of SSR minimization, that searches over the potential range of the threshold, is used to yield a consistent estimate of the threshold. The tests using the appropriate threshold estimate are conducted and the distribution of the F-statistic for testing the null of $\rho_1=\rho_2=0$ listed in Table 6.7, 6.8, 6.9, or 6.10 is employed.

Test results

The tests for the proposition of the stationarity of term spreads are reported for the three sample periods in Table 7.1, 7.2, and 7.3. The Dickey-Fuller (DF) (1979) tests for one unit root of term spread are presented in the third row, while the Enders-Granger TAR and M-TAR specifications are in the first and the second row of the tables, respectively. The null hypothesis that a spread is of the I(1) process can also be interpreted that the associated spread vector is not cointegrating.

Table 7.1. Unit root test for gnsk2-gnsk1 and gnsk3-gnsk1
(sample period = 8:1:80 ~ 7:5:85; n = 258)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(gnsk2-gnsk1)							
TAR	1	-0.168	-0.354	-914.03/	24.51	6.1	3.26
		(-3.10)[e]	(-6.42)[f]	-903.43		(0.01)	(0.51)
MTAR	1	-0.147	-0.319	-912.73/	23.74	4.79	3.23
		(-2.29)	(-6.62)	-902.13		(0.03)	(0.52)
D-F	1	-0.259		-909.93/	N/A		3.08
		(-6.49)		-902.86			(0.54)
TAR-C[g]	1	-0.165	-0.366	-915.05/	25.11	7.14	3.42
a_0 =(-0.00861)		(-3.12)	(-6.51)	-904.45		(0.01)	(0.49)
(gnsk3-gnsk1)							
TAR	1	-0.169	-0.287	-809.47/	23.19	2.99	2.42
		(-3.37)[e]	(-6.04)[f]	-798.84		(0.08)	(0.66)
MTAR	1	-0.144	-0.289	810.92/	24.04	4.45	2.20
		(-2.67)	(-6.51)	-800.28		(0.04)	(0.70)
D-F	1	-0.232		-808.46/	N/A		2.45
		(-6.56)		-801.37			(0.65)
MTAR-C[g]	1	-0.136	-0.297	-812.04/	24.69	5.57	2.16
a_0= (-0.000115)		(-2.55)	(-6.65)	-801.40		(0.02)	(0.71)

NOTE: gnsk1, gnsk2, and gnsk3 denote the 1-, 2-, and 3-month Gensaki rate, respectively.

[a] AIC = T*ln(residual sum of squares) + 2*n ; SBC = T*ln(residual sum of squares) + n*ln(T), where n = number of regressors and T = number of usable observations.

[b] Entries in this column are the sample F-statistics for testing the null of $\rho_1=\rho_2=0$.

[c] Entries in this column are the sample F-statistics for the null hypothesis that adjustments are symmetric. The corresponding significance levels are contained in brackets.

[d] Q(4) is the Ljung-Box Q-statistic for the joint hypotheses of no serial correlation among the first 4 residuals.

[e] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_1=0$.

[f] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_2=0$.

[g] The C denotes the model estimation using a consistent threshold estimate.

Table 7.2. Unit root test for cd3-cd1 and Euro1-EuroSpot, and Euro3-EuroSpot and Euro3-Euro1 (sample period = 7:12:85 ~ 1:29:93, n = 395)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(cd3-cd1)							
TAR	4	-0.120	-0.088	-1068.13/	11.01	0.69	0.83
		(-4.06)[e]	(-2.99)[f]	-1044.34		(0.41)	(0.93)
MTAR	4	-0.076	-0.144	-1070.58/	12.29	3.11	1.03
		(-2.77)	(-4.51)	-1046.78		(0.08)	(0.91)
D-F	4	-0.104		-1069.43	N/A		0.78
		(-4.62)		-1049.60			(0.94)
MTAR-C[g]	4	-0.062	-0.159	-1073.91/	14.04	6.44	1.37
a_0 = (0.0000179)		(-2.24)	(-5.10)	-1050.12		(0.01)	(0.85)
(Euro1 - EuroSpot)							
TAR	3	-0.215	-0.349	-195.59/	14.50	2.75	1.72
		(-3.51)[e]	(-4.87)[f]	-175.76		(0.10)	(0.79)
MTAR	3	-0.143	-0.393	-202.31/	18.10	9.49	2.02
		(-2.17)	(-5.97)	-182.48		(0.00)	(0.73)
D-F	3	-0.268		-194.81/	N/A		1.46
		(-5.11)		-178.95			(0.83)
MTAR-C[g]	3	-0.162	-0.625	-220.21/	28.00	28.02	1.79
a_0 = (-0.02377)		(-2.98)	(-7.14)	-220.38		(0.00)	(0.77)

Table 7.2. (Continued)

Model	lag	ρ_1	ρ_2	AIC/SBC [a]	Φ [b]	$\rho_1=\rho_2$ [c]	Q(4) [d]
(Euro3-EuroSpot)							
TAR	4	-0.110	-0.109	-123.79/	5.45	0.0002	5.52
		(-2.60) [e]	(-2.37) [f]	-99.99		(0.99)	(0.24)
MTAR	4	-0.047	-0.168	-128.05/	7.62	4.22	6.13
		(-1.04)	(-3.85)	-104.26		(0.04)	(0.19)
D-F	4	-0.109		-125.79/	N/A		5.52
		(-3.11)		-105.96			(0.24)
MTAR-C [g]	4	-0.063	-0.311	-135.85/	11.65	12.06	7.28
a_0 = (-0.03826)		(-1.80)	(-4.67)	-112.05		(0.00)	(0.12)
(Euro3-Euro1)							
TAR	1	-0.163	-0.160	-697.53/	21.72	0.004	0.70
		(-5.06) [e]	(-4.39) [f]	-685.66		(0.95)	(0.95)
MTAR	1	-0.129	-0.195	-699.43/	22.78	1.90	1.14
		(-3.79)	(-5.70)	-687.57		(0.16)	(0.89)
D-F	1	-0.162		-699.53/	N/A		0.70
		(-6.60)		-691.62			(0.95)
MTAR-C [g]	1	-0.122	-0.257	-704.36/	25.53	6.84	1.18
a_0 = (-0.00998)		(-4.27)	(-5.86)	-692.50		(0.01)	(0.88)

NOTE: cd1 and cd3 denote the 1- and 3-month CD rates, respectively; EuroSpot, Euro1, and Euro3 denote the spot, 1-, and 3-month Euroyen deposit rates, respectively.

[a] AIC = T*ln(residual sum of squares) + 2*n ; SBC = T*ln(residual sum of squares) + n*ln(T), where n = number of regressors and T = number of usable observations.

[b] Entries in this column are the sample F-statistics for testing the null of $\rho_1=\rho_2=0$.

[c] Entries in this column are the sample F-statistics for the null hypothesis that adjustments are symmetric. The corresponding significance levels are contained in brackets.

[d] Q(4) is the Ljung-Box Q-statistic for the joint hypotheses of no serial correlation among the first 4 residuals.

[e] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_1=0$.

[f] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_2=0$.

[g] The C denotes the model estimation using a consistent threshold estimate.

Table 7.3. Unit root test for cd3-cd1 and Euro1-EuroSpot, and Euro3-EuroSpot and Euro3-Euro1 (sample period = 2:5:93 ~ 10:16:98, n = 298)

Model	lag	ρ_1	ρ_2	AIC/SBC[d]	Φ[e]	$\rho_1=\rho_2$[f]	Q(4)[g]
(cd3-cd1)							
TAR	4	-0.131	-0.200	-184.31/	9.39	1.26	0.77
		(-3.12)[a]	(-3.63)[b]	-162.23		(0.26)	(0.94)
MTAR	4	-0.070	-0.275	-195.07/	15.11	12.05	0.99
		(-1.60)	(-5.48)	-172.99		(0.00)	(0.91)
D-F	4	-0.154		-185.03	N/A		0.77
		(-4.18)		-166.63			(0.94)
MTAR-C[g]	4	-0.058	-0.429	-216.05/	26.88	34.24	2.02
a_0 = (-0.01216)		(-1.50)	(-7.33)	-193.97		(0.00)	(0.73)
(Euro1-EuroSpot)							
TAR	1	-0.396	-0.478	657.31/	30.28	0.41	5.04
		(-3.30)[a]	(-7.43)[b]	668.50		(0.52)	(0.28)
MTAR	1	-0.271	-0.522	653.43/	32.60	4.28	4.89
		(-2.45)	(-7.93)	664.62		(0.04)	(0.30)
D-F	1	-0.463		655.72/	N/A		4.89
		(-7.76)		663.19			(0.30)
MTAR-C[g]	1	-0.242	-0.547	650.85/	34.15	6.88	5.10
a_0 = (-0.03033)		(-2.35)	(-8.14)	662.04		(0.01)	(0.28)

Table 7.3. (Continued)

Model	lag	ρ_1	ρ_2	AIC/SBC[d]	Φ[e]	$\rho_1=\rho_2$[f]	Q(4)[g]
(Euro3-EuroSpot)							
TAR	1	-0.293	-0.219	593.19/	15.61	0.69	5.66
		(-3.81)[a]	(-4.37)[b]	604.21		(0.41)	(0.23)
MTAR	1	-0.339	-0.179	590.12/	17.30	3.76	5.75
		(-5.06)	(-3.34)	601.14		(0.05)	(0.22)
D-F	1	-0.240		591.89/	N/A		5.73
		(-5.53)		599.23			(0.22)
MTAR-C[g]	1	-0.461	-0.165	583.23/	21.17	10.74	5.75
a_0 = (0.06477)		(-5.78)	(-3.39)	594.25		(0.00)	(0.22)
(Euro3-Euro1)							
TAR	1	-0.239	-0.327	369.27/	24.42	1.13	6.36
		(-3.49)[a]	(-6.29)[b]	380.29		(0.29)	(0.17)
MTAR	1	-0.445	-0.188	360.20/	29.75	10.28	5.25
		(-7.08)	(-3.49)	371.22		(0.00)	(0.26)
D-F	1	-0.296		368.41/	N/A		6.47
		(-6.91)		375.76			(0.17)
MTAR-C[g]	1	-0.497	-0.173	354.52/	33.18	16.17	5.16
a_0 = (0.02412)		(-7.62)	(-3.34)	365.54		(0.00)	(0.27)

NOTE: cd1 and cd3 denote the 1- and 3-month CD rates, respectively; EuroSpot, Euro1, and Euro3 denote the spot, 1-, and 3-month Euroyen deposit rates, respectively.

[a] AIC = T*ln(residual sum of squares) + 2*n ; SBC = T*ln(residual sum of squares) + n*ln(T), where n = number of regressors and T = number of usable observations.

[b] Entries in this column are the sample F-statistics for testing the null of $\rho_1=\rho_2=0$.

[c] Entries in this column are the sample F-statistics for the null hypothesis that adjustments are symmetric. The corresponding significance levels are contained in brackets.

[d] Q(4) is the Ljung-Box Q-statistic for the joint hypotheses of no serial correlation among the first 4 residuals.

[e] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_1=0$.

[f] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_2=0$.

[g] The C denotes the model estimation using a consistent threshold estimate.

The empirical evidence can be summarized as follows:

A. The first sample period:

1. The residual sequence of the equilibrium relationship, as obtained from the first step of test procedure, is estimated under the structure of the TAR and M-TAR specifications using the threshold $a_0=0$, and also in the form of the Dickey-Fuller specification. Diagnostic statistics as well as AIC and SBC values select one lag of first difference of the data series to be included into the model specifications for the term spread of 2- and 1-month gensaki rates, and the spread of 3- and 1-month gensaki rates. The optimal lag lengths are very consistent across the three test models.

2. The unit root hypothesis: $\rho_1=\rho_2=0$ is rejected for both of the yield spreads, by comparing the sample values to the appropriate test statistics Φ_μ and Φ_μ^* corresponding to TAR and M-TAR model, at the 1% significance level of 6.47 and 6.99, respectively.

3. Given that the stationality of the term spread is concluded, the subsequent standard F-tests of the null hypothesis of symmetric adjustment: $\rho_1=\rho_2$ are conducted. The test results reject the symmetry hypothesis and, thus, conclude asymmetric adjustments in the process of yield spreads, at the 5% significance level.

4. The sample value of τ in the Dickey-Fuller model also indicates that the spread is stationary. The estimate of the adjustment parameter of ρ lies between the estimate of ρ_1 and ρ_2.

5. The best fitting asymmetric model is chosen by measuring AIC/SBC. Since a priori knowledge about the true value of a_0 is limited, Chan's (1993) method is employed to find the

consistent estimate of the threshold. For the analysis of 2- and 1-month gensaki rate differential, a consistent threshold estimate of -0.00861 in the framework of TAR adjustments is obtained by Chan's procedure, given that the TAR specification performs slightly better than the M-TAR and the linear attractor specifications under the AIC/SBC rule. A consistent threshold estimate of -0.000115 is selected for the spread between 3- and 1-month gensaki rates, with a M-TAR adjustment process.

6. The point estimates of ρ_1 and ρ_2 suggest convergence in a fashion that the speed of adjustment depends on the state of discrepancies from the threshold a_0. The sample statistics obtained from the estimation utilizing the appropriate estimate of the threshold are compared with corresponding test statistics $\Phi_\mu(c)$ and $\Phi_\mu^*(c)$. Likewise, the sample value $\Phi_\mu(c)$ of 25.11 and the sample statistic $\Phi_\mu^*(c)$ of 24.69 for the spread between 2- and 1-month gensaki rates and between 3- and 1-month gensaki rates, allows the null hypothesis of a unit root to be rejected at the 1% significance level. The F-test for symmetric adjustment rejects the null of symmetry at the 1% significance level as well.

7. As measured by AIC and SBC, the selected asymmetric model which incorporates the consistent threshold estimator fits the data substantially better than the other models. Overall, the asymmetric specification yields a smaller AIC/SBC than the symmetric model.

8. The results indicate that the yield spreads are stationary and the adjustment mechanisms are asymmetric. The estimates conclude threshold adjustment in the spread of 2- and 1-month gensaki rates, in which the attractor is stronger when the spread is below the threshold than above. On the other hand, the estimates infer the momentum threshold process for the 3- and 1- month spread, in which the autoregressive decay is relatively sharp when it is

decreasing rather than when it is increasing.

B. The second sample period:

1. The test statistics reject the null hypothesis that the spreads contain a unit root and conclude the stationarity of the spreads on this basis.

2. M-TAR is chosen over the TAR specification as the best fitting model. Moreover, the M-TAR model test with the consistent threshold estimator is more appropriate than the Dickey-Fuller test under the model selection criteria of AIC/SBC.

3. The test results reject the null hypothesis of symmetry: $\rho_1 = \rho_2$ at the 1% significance level. The estimates indicate the asymmetric dynamic process of the spreads, which is characterized by a relatively greater extent of momentum for adjustment in the downward direction rather than upward.

C. The third sample period:

1. Much the same as occurs in the first and the second sample periods. The third sample tests support the expectations hypothesis of the interest rate term structure. As measured by AIC/SBC, the M-TAR model with a consistent estimate of the threshold explains the data to a better extent than the Dickey-Fuller or the TAR specifications. Diagnostic checking of the residuals indicates appropriate lag changes included in the model. The F-test for symmetric adjustment is rejected at the 1% significance level. Thus, the empirical evidence implies stationary interest rate differentials with the asymmetric adjustment mechanism.

2. The point estimates for spread ρ_1 and ρ_2 suggest faster convergence for the negative rather than positive deviation of yield spread changes from the threshold, for 3- and 1-month CD,

and for 1-month and spot Euroyen deposit rate differentials. In contrast, for the series of yield spreads between 3- and 1-month as well as between 3-month and spot Euroyen deposits rates, the speed of adjustment is more rapid for positive than for negative discrepancies from the threshold.

The open-market operations of the BOJ can be categorized as defensive ones directed toward offsetting the effects of short-run temporary fluctuations in the demand for high-power money in money markets, and the dynamic ones aimed at changing the level of the target.

The BOJ has long targeted interbank interest rates in its daily and short-term operations. The analysis for each of the three sample periods suggests that the positive term spreads resulting from temporary falls in daily or very short-term rates from the persistent target rate are eliminated relatively swiftly whereas the alternative movements display persistence. In the presence of transitory decreases in interbank and money market rates, the daily or short-term operations by the BOJ are relatively heavy and the subsequent daily rates are expected to increase and move quickly back to the target level. On the other hand, the operations are less intensive to accommodate the transitory shortage of funds in money markets, which lead to slow interest rates adjustments to the target rate.

For the first and the second periods, the estimates of the speed of adjustment terms in the analysis of yield spreads between relatively medium-term rates indicated that increases are moderate and prolonged whereas decreases revert swiftly toward the attractor in the adjustment paths. It might be implied that increase in the target level of the Bank is attained by sequential target adjustments over time, while a decrease is achieved by sharp sizable changes over a short horizon.

For the third sample period, nevertheless, the estimates suggest that the rises are quickly altered by subsequent falls but the falls display a large amount of persistence. The asymmetric dynamic process in interest rate series might be attributed to the long recession in Japan after the collapse of the asset-price bubble.

The cointegration test of the term structure are reported in Table 7.4, 7.5, and 7.6. The third row of the tables refers to the Engle-Granger (EG) tests, and the first and the second rows are the Enders-Siklos TAR and M-TAR model tests, respectively.

The tests support the prediction under the expectations hypothesis that the rates of different maturities are cointegrated. The appropriate asymmetric test methodology selected via the AIC/SBC rule performs better than the Engle-Granger test model. Diagnostic checking of the residuals of the fitted models shows no evidence of serial correlation. The tests for the equal adjustment coefficient: $\rho_1 = \rho_2$ are rejected with F values significant at the 0.00 level. For each of the three periods, the conclusion based on the cointegration test results is remarkably similar to that of the unit-root test. The exception lies merely on the model specification in the analysis of the spreads, between 3- and 1-month gensaki rates over the first time interval, in which the cointegration test suggests a TAR whereas the unit-root test concludes an M-TAR representation.

Table 7.4. Cointegration test for gnsk2, gnsk1 and gnsk3, gnsk1
(sample period = 8:1:80 ~ 7:5:85, n = 258)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(gnsk2, gnsk1, constant)							
TAR	1	-0.149	-0.348	-931.21/	22.76	7.19	3.36
		(-2.91)[e]	(-6.21)[f]	-920.61		(0.01)	(0.50)
MTAR	1	-0.134	-0.302	-928.93/	21.43	4.87	3.24
		(-2.19)	(-6.28)	-918.33		(0.03)	(0.52)
E-G	1	-0.239		-926.04/	N/A		3.24
		(-6.12)		-918.97			(0.52)
TAR-C[g]	1	-0.145	-0.364	-932.68/	23.62	8.69	3.35
a_0 = (-0.00825)		(-2.91)	(-6.35)	-922.08		(0.00)	(0.50)
(gnsk3, gnsk1, constant)							
TAR	1	-0.146	-0.299	-830.59/	22.11	5.3	2.52
		(-3.14)[e]	(-5.99)[f]	-819.99		(0.02)	(0.64)
MTAR	1	-0.128	-0.275	-830.01/	21.77	4.72	2.02
		(-2.40)	(-6.25)	-819.41		(0.03)	(0.73)
E-G	1	-0.217		-827.28/	N/A		2.50
		(-6.19)		-820.21			(0.64)
TAR-C[g]	1	-0.138	-0.315	-832.19	23.04	6.92	2.42
a_0 = (-0.00885)		-3.03	(-6.18)	-821.59		(0.01)	(0.66)

NOTE: gnsk1, gnsk2, and gnsk3 denote the 1- , 2-, and 3-month Gensaki rate, respectively.

[a] AIC = T*ln(residual sum of squares) + 2*n ; SBC = T*ln(residual sum of squares) + n*ln(T), where n = number of regressors and T = number of usable observations.

[b] Entries in this column are the sample F-statistics for testing the null of $\rho_1=\rho_2=0$.

[c] Entries in this column are the sample F-statistics for the null hypothesis that adjustments are symmetric. The corresponding significance levels are contained in brackets.

[d] Q(4) is the Ljung-Box Q-statistic for the joint hypotheses of no serial correlation among the first 4 residuals.

[e] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_1=0$.

[f] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_2=0$.

[g] The C denotes the model estimation using a consistent threshold estimate.

Table 7.5. Cointegration test for cd3, cd1 and Euro1, EuroSpot, and Euro3, EuroSpot and Euro3, Euro1 (sample period = 7:12:85 ~ 1:29:93, n = 395)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(cd3, cd1, constant)							
TAR	4	-0.119	-0.086	-1072.45/	10.90	0.77	0.84
		(-4.04)[e]	(-2.96)[f]	-1048.66		(0.38)	(0.93)
MTAR	4	-0.075	-0.142	-1074.79/	12.12	3.08	1.03
		(-2.74)	(-4.48)	-1050.99		(0.08)	(0.91)
E-G	4	-0.103		-1073.68	N/A		0.78
		(-4.59)		-1053.84			(0.94)
MTAR-C[g]	4	-0.061	-0.157	-1078.01/	13.81	6.28	1.35
a_0 = (0.0000152)		(-2.22)	(-5.06)	-1054.21		(0.01)	(0.85)
(Euro1, EuroSpot, constant)							
TAR	3	-0.224	-0.351	-230.84/	15.36	2.45	1.10
		(-3.41)[e]	(-5.13)[f]	-211.01		(0.12)	(0.89)
MTAR	3	-0.137	-0.421	-240.40/	20.52	12.07	1.30
		(-2.03)	(-6.39)	-220.57		(0.00)	(0.86)
E-G	3	-0.284		-230.37/	N/A		0.88
		(-5.31)		-214.50			(0.93)
MTAR-C[g]	3	-0.166	-0.636	-257.45/	30.04	29.81	1.55
a_0 = (-0.01836)		(-2.98)	(-7.70)	-237.62		(0.00)	(0.82)

Table 7.5. (Continued)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(Euro3, EuroSpot, constant)							
TAR	5	-0.118	-0.126	-200.02/	7.04	0.02	6.99
		(-2.73)[e]	(-2.96)[f]	-172.26		(0.88)	(0.14)
MTAR	5	-0.047	-0.192	-206.72/	10.47	6.64	7.11
		(-1.07)	(-4.55)	-178.95		(0.01)	(0.13)
E-G	5	-0.122		-202.00/	N/A		6.96
		(-3.76)		-178.20			(0.14)
MTAR-C[g]	5	-0.038	-0.266	-216.11/	15.39	16.11	8.46
a_0 = (-0.00859)		(-1.01)	(-5.54)	-188.35		(0.00)	(0.08)
(Euro3, Euro1, constant)							
TAR	3	-0.162	-0.151	-722.24	17.23	0.06	0.21
		(-4.70)[e]	(-4.10)[f]	-702.41		(0.81)	(1.00)
MTAR	3	-0.123	-0.192	-724.33/	18.36	2.12	0.34
		(-3.46)	(-5.35)	-704.50		(0.15)	(0.99)
E-G	3	-0.157		-724.18/	N/A		0.21
		(-5.87)		-708.32			(0.99)
MTAR-C[g]	3	-0.113	-0.275	-731.91/	22.49	9.72	0.42
a_0 = (-0.00940)		(-3.81)	(-5.96)	-712.08		(0.00)	(0.98)

NOTE: cd1 and cd3 denote the 1- and 3-month CD rates, respectively; EuroSpot, Euro1, and Euro3 denote the spot, 1-, and 3-month Euroyen deposit rates, respectively.

[a] AIC = T*ln(residual sum of squares) + 2*n ; SBC = T*ln(residual sum of squares) + n*ln(T), where n = number of regressors and T = number of usable observations.
[b] Entries in this column are the sample F-statistics for testing the null of $\rho_1=\rho_2=0$.
[c] Entries in this column are the sample F-statistics for the null hypothesis that adjustments are symmetric. The corresponding significance levels are contained in brackets.
[d] Q(4) is the Ljung-Box Q-statistic for the joint hypotheses of no serial correlation among the first 4 residuals.
[e] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_1=0$.
[f] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_2=0$.
[g] The C denotes the model estimation using a consistent threshold estimate.

Table 7.6. Cointegration test for cd3, cd1 and Euro1, EuroSpot, and Euro3, EuroSpot and Euro3, Euro1 (sample period = 2:5:93 ~ 10:16:98, n = 298)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(cd3, cd1, constant)							
TAR	4	-0.192	-0.232	-216.46/	12.91	0.41	0.86
		(-3.92)[e]	(-4.16)[f]	-194.38		(0.52)	(0.93)
MTAR	4	-0.113	-0.346	-229.67/	20.12	13.66	0.95
		(-2.36)	(-6.30)	-207.58		(0.00)	(0.92)
E-G	4	-0.208		-218.04	N/A		0.86
		(-5.05)		-199.64			(0.93)
MTAR-C[g]	4	-0.105	-0.444	-242.02/	27.17	26.61	1.75
a_0 = (-0.01225)		(-2.36)	(-7.35)	-219.94		(0.00)	(0.78)
(Euro1, EuroSpot, constant)							
TAR	1	-0.376	-0.508	657.55/	31.79	1.17	4.84
		(-3.37)[e]	(-7.62)[f]	668.74		(0.28)	(0.30)
MTAR	1	-0.278	-0.545	653.64/	34.15	5.08	4.45
		(-2.60)	(-8.12)	664.83		(0.02)	(0.34)
E-G	1	-0.478		656.73/	N/A		4.65
		(-7.90)		664.19			(0.33)
MTAR-C[g]	1	-0.265	-0.583	650.59/	36.01	8.17	4.83
a_0 = (-0.06806)		(-2.78)	(-8.30)	661.78		(0.00)	(0.31)

Table 7.6. (Continued)

Model	lag	ρ_1	ρ_2	AIC/SBC[a]	Φ[b]	$\rho_1=\rho_2$[c]	Q(4)[d]
(Euro3, EuroSpot, constant)							
TAR	2	-0.301	-0.262	642.47/	18.41	0.21	0.50
		(-4.35)[e]	(-4.64)[f]	657.26		(0.64)	(0.97)
MTAR	2	-0.362	-0.213	639.42/	20.11	3.24	0.54
		(-5.51)	(-3.68)	654.21		(0.07)	(0.97)
E-G	2	-0.277		640.69/	N/A		0.52
		(-6.06)		651.78			(0.97)
MTAR-C[g]	2	-0.523	-0.199	630.55/	25.17	12.23	0.47
a_0 = (0.07847)		(-6.27)	(-3.96)	-645.33		(0.00)	(0.98)
(Euro3, Euro1, constant)							
TAR	2	-0.200	-0.338	387.47/	21.64	3.05	4.06
		(-3.23)[e]	(-6.04)[f]	402.25		(0.08)	(0.40)
MTAR	2	-0.401	-0.184	382.96/	24.21	7.57	4.00
		(-6.41)	(-3.36)	397.75		(0.01)	(0.41)
E-G	2	-0.277		388.54/	N/A		4.18
		(-6.32)		399.63			(0.38)
MTAR-C[g]	2	-0.460	-0.171	377.53/	27.36	13.12	4.32
a_0 = (0.03232)		(-6.93)	(-3.31)	392.32		(0.00)	(0.36)

NOTE: cd1 and cd3 denote the 1- and 3-month CD rates, respectively; EuroSpot, Euro1, and Euro3 denote the spot, 1-, and 3-month Euroyen deposit rates, respectively.

[a] AIC = T*ln(residual sum of squares) + 2*n ; SBC = T*ln(residual sum of squares) + n*ln(T), where n = number of regressors and T = number of usable observations.
[b] Entries in this column are the sample F-statistics for testing the null of $\rho_1=\rho_2=0$.
[c] Entries in this column are the sample F-statistics for the null hypothesis that adjustments are symmetric. The corresponding significance levels are contained in brackets.
[d] Q(4) is the Ljung-Box Q-statistic for the joint hypotheses of no serial correlation among the first 4 residuals.
[e] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_1=0$.
[f] Entries in the brackets of this column are the t-statistic for the null hypothesis $\rho_2=0$.
[g] The C denotes the model estimation using a consistent threshold estimate.

Error-Correction Representation

The positive finding of an equilibrium relationship with asymmetric adjustment justifies the estimation of the asymmetric error-correction representation. The consistent estimate of the threshold for each of the equations in the error-correction model is obtained by Chan's procedure. The most appropriate lag length is selected by the measure of the multivariate AIC/SBC. Using consistent estimates of the threshold, the estimated error-correction equations for the term spread between daily and 3-month, and 1- and 3-month Euroyen deposit rates are as follows:

<u>Euro3-EuroSpot Case</u> (Sample period: 2:3:1993 ~ 10:16:1998)

$$\Delta EuroSpot_t = \underset{(-1.54)}{-0.012}\ \underset{(-2.37)}{-0.149}\Delta EuroSpot_{t-1} \underset{(0.66)}{+0.036}\Delta Euro3_{t-1} \underset{(5.28)}{+0.350}z\text{-}puls_{t-1} \underset{(1.02)}{+0.042}z\text{-}minus_{t-1}$$

$$\Delta Euro3_t = \underset{(-1.30)}{-0.013}\ \underset{(-0.59)}{-0.045}\Delta EuroSpot_{t-1} \underset{(0.62)}{+0.043}\Delta Euro3_{t-1} \underset{(-0.89)}{-0.057}z\text{-}plus_{t-1} \underset{(-3.03)}{-0.180}z\text{-}minus_{t-1}$$

where: $z\text{-}plus_{t-1} = I_t(Euro3_{t-1} - EuroSpot_{t-1} + 0.021)$,

$z\text{-}minus_{t-1} = (1 - I_t)(Euro3_{t-1} - EuroSpot_{t-1} + 0.021)$.

<u>Euro3-Euro1 Case</u> (Sample period: 2:3:1993 ~ 10:16:98)

$$\Delta Euro1_t = \underset{(-2.14)}{-0.023}\ \underset{(-5.48)}{-0.507}\Delta Euro1_{t-1} \underset{(3.33)}{+0.384}\Delta Euro3_{t-1} \underset{(7.91)}{+0.782}z\text{-}plus_{t-1} \underset{(-0.34)}{-0.030}z\text{-}minus_{t-1}$$

$$\Delta Euro3_t = \underset{(-2.10)}{-0.018}\ \underset{(-8.06)}{-0.626}\Delta Euro1_{t-1} \underset{(6.44)}{+0.624}\Delta Euro3_{t-1} \underset{(4.15)}{+0.343}z\text{-}plus_{t-1} \underset{(-3.47)}{-0.261}z\text{-}minus_{t-1}$$

where: $z\text{-}plus_{t-1} = I_t(Euro3_{t-1} - Euro1_{t-1} + 0.00276)$,

$z\text{-}minus_{t-1} = (1 - I_t)(Euro3_{t-1} - Euro1_{t-1} + 0.00276)$,

and: I_t = momentum threshold Heaviside indicator function, of which the unbiased threshold estimate is obtained by applying Chan's (1993) method to each equation.

The key feature in the fitted equations is the pattern of the estimated coefficients for *z-plus* and *z-minus*. The point estimates of the error-correction terms in the Euro3-EuroSpot case indicate that the Euroyen spot rate adjusts strongly to positive changes in deviations from a long-run equilibrium and moderately to negative changes, while the 3-month rate adjusts to negative, but not positive changes in the discrepancies.

The spot rate rises to eliminate approximately 35% of a one-unit deviation from the long-run equilibrium level when the gap is increasing, and about 4% of the error correction term when the gap is decreasing. On the other hand, the 3-month rate is estimated to decline by 18% of the error term when the gap is decreasing, but the rate is not responsive when the gap is increasing. Thus, the negative changes in the discrepancies from the long-run equilibrium relationship stemming from the decline of the 3-month rate or the rise of the spot rate tend to be prolonged, as the discrepancies are eliminated relatively more gradually than changes in the other direction.

In the Euro3-Euro1 case, the estimation for the term spread between the 3- and 1-month rate also indicate to some extent sluggish adjustment to the long-run equilibrium when the error correction term is decreasing rather than when it is increasing. Nevertheless, the 1-month rate responds to positive, but not negative, changes in the deviations of the spread from its long-run level, whereas the 3-month rate responds to positive as well as negative changes in the gap. Note that the 3-month rate adjusts in the "wrong" direction in response to a increase in the gap.

In contrast, the error-correction models assuming symmetric adjustment over the sample period of 2:5:1993-10:16:1998 are:

Euro3-EuroSpot Case

$$\underset{}{\Delta EuroSpot_t} = \underset{(-1.29)}{-0.010} \underset{(-3.48)}{-0.217}\Delta EuroSpot_{t-1} \underset{(0.02)}{+0.001}\Delta Euro3_{t-1} \underset{(3.44)}{+0.126}\hat{\mu}_{t-1}$$

$$\Delta Euro3_t = \underset{(-0.95)}{-0.009} \underset{(-0.81)}{-0.062}\Delta EuroSpot_{t-1} \underset{(0.29)}{+0.020}\Delta Euro3_{t-1} \underset{(-2.77)}{-0.124}\hat{\mu}_{t-1}$$

where: $\hat{\mu}_{t-1} = (Euro3_t - EuroSpot_t + 0.021)$.

Euro3-Euro1 Case

$$\Delta Euro1_t = \underset{(-1.10)}{-0.012} \underset{(-4.57)}{-0.451}\Delta Euro1_{t-1} \underset{(2.08)}{+0.252}\Delta Euro3_{t-1} \underset{(4.58)}{+0.333}\hat{\mu}_{t-1}$$

$$\Delta Euro3_t = \underset{(-1.15)}{-0.010} \underset{(-7.13)}{-0.583}\Delta Euro1_{t-1} \underset{(5.21)}{+0.524}\Delta Euro3_{t-1} \underset{(0.19)}{+0.011}\hat{\mu}_{t-1}$$

where: $\hat{\mu}_t = (Euro3_t - Euro1_t + 0.00276)$.

In the system equations of the Euro3-Euro1 case, the t-statistic for the error correction term on the 3-month rate is insignificant at conventional levels. As such, the symmetric adjustment model implies that the 3-month rate is weakly exogenous, and only the 1-month rate adjusts to eliminate deviations from the long-run equilibrium property. Even though the threshold representation involves additional coefficients in the model equations, the multivariate values of AIC and SBC suggest the M-TAR adjustment over the symmetric specification. The multivariate AIC and SBC are -2303.74 and -2266.90, respectively, for the consistent threshold M-TAR model and -2291.93 and -2262.46, respectively, for the symmetric dynamics of the spread between 3-month and spot rate, and the values are -2443.00 and -2406.16 for the system specified in M-TAR adjustments and -2409.15 and -2379.68 in symmetric adjustments for the spread between 3- and 1-month rate.

Impulse response analysis

The impulse response analysis, which examines the percentage response of variables over periods to an initial one-unit-standard deviation shock to each of the variables in the system, can provide further insight into the economic dynamics. The traditional impulse response function for the level of the system variables takes the form in which the effect of a shock on the behavior of time series is symmetric, linear, and history-dependent. As such, a negative shock has precisely the opposite impact of a positive shock. A two-unit shock imposes exactly twice the impact than a one-unit shock. In addition, the past realizations of system variables do not affect the impulse response.

Nevertheless, the traditional impulse response function is less effectively employed to the nonlinear model than to linear ones. It was shown by Koop, Pesearson, and Potter (1996) that nonlinear models produce impulse responses that are contingent on the magnitude and sign of the shock chosen by the researcher, and also rely on the history specifically selected as the basis for comparison of the states of the two systems, where one is subject to a one-time initial shock while the other is not. Hence, the traditional impulse response analysis, which does not take into account the joint, conditional, and marginal densities of the impulse response function as a random vector based on the probability space of the time series of system variable, cannot be usefully applied to the asymmetric models under consideration in this research. As such, an impulse response analysis was not conducted in this study.

CHAPTER 8. CONCLUSIONS

This research focused on the behavior of the term structure of interest rates in Japan during the 1980s and 1990s. A nonlinear approach in testing the expectations hypothesis of the term structure was then proposed. The analysis was triggered by the theoretical advance recently introduced by Enders and Granger (1998), and Enders and Siklos (1998), which generalizes the conventional Dickey-Fuller (1979) and Engle-Granger (1987) procedures to allow TAR and Momentum-TAR adjustment mechanisms in the dynamic processes of time series.

It is generally agreed that interest rates are integrated of order 1. Nevertheless, the implicit assumption of symmetric adjustment is problematic since recent empirical studies and the observed behavior of interest rates have implied generally that the dynamic adjustments of yield spreads might be different in the presence of positive deviations versus negative deviations from the long run equilibrium. Balke and Fomby (1997), and Enders and Granger (1998) also confirmed asymmetric adjustment in the term structure of US interest rates.

This research presents a detailed discussion on the monetary policies of Japan and asymmetric adjustments in term structure. By taking into account the asymmetry in term structure, this research analyzed the Japanese gensaki, CD, and Euroyen deposit rates and found empirical support for the expectations hypothesis, in the sense that the yield spread can be consistently justified by rational forecasts of future movements in short-term interest rates. The asymmetric error-correction approach was employed to characterize the dynamics of interest rate differentials as well. The estimation showed the asymmetric adjustments of interest rate term structure of Japan, which, accompanied by the maintained expectations hypothesis, can be explained by the manner in which the Bank of Japan controls the interbank rate.

A conventional impulse response analysis following the error-correction estimation of yield spread would be inappropriate, since impulse response generated by nonlinear models would be contingent on history and shock (Potter,1994). Furthermore, the shocks to different variables in model system might be of compositional dependence in multivariate models (Lee & Pesaran, 1993). Thus, a more appropriate approach for future study would be an impulse response function generalized to measure the persistence and the asymmetric effects of shocks.

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